# THREE MACROECONOMIC ESSAYS ON THE UK ECONOMY

DISSERTATION

A dissertation submitted in partial fulfillment of the requirements for the degree of Doctor of Philosophy in the Department of Economics at the University of Sheffield

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Sheffield, June 2020

# ABSTRACT OF DISSERTATION

# THREE MACROECONOMIC ESSAYS ON THE UK ECONOMY

This dissertation consists of three essays that make contributions to the research on the empirical analysis of macroeconomic issues of the UK economy. The first essay elaborates on the determinants of asset price co-movements and examines the role of monetary variables in predicting changes in the dynamic co-movements between stock and house markets returns. The second essay identifies changes in the historical growth rates of real output, understands the asymmetric behavior over real output expansions and contractions, and considers whether the growth rate of the debt-to-GDP ratio is a leading indicator of the time-varying transition probabilities between expansions and contractions. The third essay focuses on trade relations between UK, four big euro area countries, and the USA, and in this global context we forecast the impact of a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to German real output, which proxies an expansionary shock in the euro area, on the economy of all countries in the analysis.

KEYWORDS: Monetary policy, dynamic conditional correlation approach, quantile approach, principal component analysis, economic growth, public debt, Markov Switching model, trade flows, Global VAR

# DECLARATION

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# DEDICATION

To my Family

# ACKNOWLEDGMENTS

Firstly, I would like to express my sincere gratitude to my supervisor Dr. Kostas Mouratidis for all the support, encouragement and inspiration he gave me to continue my research and complete my PhD thesis. I would also like to thank the participants of the departmental PhD conferences for their insight comments and suggestions which have improved my thesis. And last but not least, I would like to thank my family for their spiritual support throughout the writing of my thesis.

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# Chapter 1

# Introduction

The global financial meltdown 2007-2009 was one of the most remarkable and noteworthy events that happened during the last decade. It has had a great impact on all the major countries around the world, including the United Kingdom (UK), which is the focus of the present thesis. As a result, numerous empirical studies and papers have worked on the implications of this crisis and the events that followed. More specifically, there has been lots of research/literature regarding the examination of monetary policy effects on individual asset returns, and the influence of macroeconomic and risk factors on stock-bond correlations. One of the main concerns of the governments is the establishment of financial stability which should complement the establishment of macroeconomic goals such as real growth and stable inflation rate. Given that asset price co-movements are of great significance to policy makers and portfolio investors, the first essay of this dissertation adds to literature on the determinants of asset price co-movements by examining the role of monetary policy developments in the UK in driving changes in dynamic correlations between stock and house returns, given the contribution of other macroeconomic variables. The importance of this topic stems from the fact that, firstly stock and house markets play a crucial role in the transmission mechanism of monetary policy, and secondly stocks and houses constitute large components of the total wealth of UK households, thus affecting consumption and investment decisions. Initially, we use the dynamic conditional correlation model to estimate the covariance of stock-house returns, where an asymmetry in the correlation between the two returns has been observed. In turn, we use the method of least squares with breaks to filter the variability of dynamic conditional covariance between the stock market return and house market return. The variability constitutes the cyclical component of dynamic covariance which reflect periods of positive and negative deviations from trend. If deviations are large, boom and bust outcomes are present and imply risk. Managing risk means having information about the whole distribution of these possible outcomes. To get information about these outcomes, we relate the variability of dynamic conditional covariance with developments in monetary policy and examine whether predetermined monetary policy measures have predictive content to understanding changes in dynamic covariance of stock-house returns across the entire conditional distribution of the response variable using a quantile approach which is the case with the present research. Modeling the stylized facts of dynamic co-movements of stock-house returns, this study will provide policy makers with valuable information about monetary policy developments in characterizing these facts, and benefit investors in diversifying their portfolio strategies.

The global financial crisis has also produced disappointing economic outcomes which are conducive to economic growth. The transformation of the financial crisis to sovereign debt crisis has questioned the long-term fiscal sustainability in advanced and developing economies, given the fact that the debt-to-GDP ratios have reached high levels. A crisis might bring negative effect into the real economy not only by destroying financial wealth and obstructing financial intermediation, but also by aggravating fiscal positions that brings public debt beyond certain levels which may not be sustainable. In the aftermath of the financial crisis, the importance of fiscal sustainability for long-term economic performance has gained a new research interest. Several empirical studies have examined the link between public debt and real output growth, without capturing an empirical fact that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods. This empirical regularity cannot be explained by linear models used in the literature which cannot capture these distinct facts, but instead a nonlinear model capturing these distinct patterns in the data is regarded more appropriate. A nonlinear model which is suitable to analyze different dynamic structures of the data over a time period and allows switching between these structures is the notorious Markov switching model. The second essay of this dissertation adds to literature by fitting a Markov switching model to annual real output growth rate in order to examine changes in mean during three centuries. The first objective is to identify changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. The second objective is to examine whether the growth rate of the debt-to-GDP ratio is a leading indicator of the time-varying transition probabilities between expansions and contractions of the real output growth, which is true in the present case as increases in the lagged growth rate of the debt ratio are associated with a lower probability of remaining in the expansion regime, and a higher probability of moving from the high growth regime to low growth regime. The link between real output growth and debt is an important topic from a policy perspective. Modeling the stylized facts of the asymmetric behavior over real output expansions and contractions and establishing that increases in the growth rate of the debt-to-GDP ratio lead to switching from expansion to contraction regimes will support the arguments for long-run fiscal consolidation.

On January 31th, 2020, the UK withdrew from the European Union (EU) after almost half a century of membership. The process of Brexit (leaving the EU) had begun with the outcome of Britain's referendum, which took place on the 23$^{rd}$ of June 2016. The decision to leave the EU has spurred a discussion about the potential risks for the UK economy, since the EU constitutes its largest trading partner. Although some studies suggest that the potential costs of Brexit to the UK economy could be substantial, the literature is rather limited, and this requires more research to examine the significance and duration of the potential risks for the UK and the EU economies. This chapter aspires to fill in this gap in the literature by using the innovative tool of the global vector autoregression (GVAR) to model the trade interactions between the UK, the four big euro area countries – Germany, France, Italy and Spain – and the USA. Then, we use the model to perform impulse response analysis and forecast error variance decomposition. In particular, we forecast the effects of a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area, on the economy of all countries in the analysis. In addition, we compute the proportion of the forecast variance of real imports and real exports of the UK and Germany that is explained by shocks to all variables and countries in the model. This research is important to assess the potential effects of Brexit on the economy of all countries in the analysis and the duration of these effects.

All in all, UK has always been and is still one of the main countries worldwide to affect the global economy and thus with this dissertation, we address this importance by examining the UK economy from three different macroeconomic perspectives, the stock-house relationship, the debt-growth relationship and the international trade.

# Chapter 2

# The predictive content of monetary variables for the variability of dynamic covariance between stock and house returns

## 2.1. Introduction

In a modern economy with liberalized asset markets, developments in asset prices play a significant role. They affect the real economy through the wealth effect which impacts consumption, and Tobin's q and the financial accelerator effect which influence investment. Tobin's q is the ratio of the market value of firms' installed real capital to the replacement cost of this capital stock. If this ratio is greater to one, it implies that the market value of installed capital is greater than its replacement cost and thus firms have incentive to invest more in capital (Tobin, 1969). Thus, the q-ratio provides a link between the stock market and the real economy. In addition, asset prices provide information about market expectations regarding the future course of macroeconomic environment. The globalization of asset markets has facilitated the spread of financial shocks all over the world, producing unusual asset price co-movements with strong episodes of booms and busts. These asset price developments, which the world economy has experienced during the last decade of the twentieth century, have showed that the stylized facts that characterized the asset price co-movements are of great importance to policy makers and portfolio investors (Gomes and Taamouti, 2014 and 2016). For policy makers these facts are related to monetary policy and financial stability frameworks, while for investors these facts are related to portfolio diversification strategies.

The modern portfolio theory (Markowitz, 1952) offers a reasoning of asset price co-movements. Let us assume, for instance, that investors hold a portfolio of two assets, and the price of an asset increases, thus disturbing the weights of the portfolio by increasing the value of this asset. Then, investors will rebalance their portfolio by increasing the demand for the other asset, thus increasing its price. Therefore, the two asset prices move together. In addition, the traditional asset pricing model of discounted expected cash flows offers a reasoning why asset prices can move as a result of policy actions. Particularly, monetary policy impacts asset prices firstly by affecting the risk-free interest rate, which together with the risk premium constitute the discount rate (Barsky and Bogusz, 2014), and secondly by influencing the aggregate liquidity of the economy which consequently influences real output in the short-run (Borgy et al., 2009). To understand the link between monetary policy and asset prices, we write the traditional asset pricing model in a compact form known as Gordon specification:

$$P_t \frac{D_t}{r_t + \theta - g} \tag{2.1.1}$$

where $P_t$ is the real asset price, $D_t$ is the real dividend, $r_t$ is the risk-free interest rate, $\theta$ is the required compensation of risk and $g$ is the growth rate of real dividends. The parameters $\theta$ and $g$ are regarded as constant. Proceeding as in Machado and Sousa (2006), we assume that real dividends are positively related with real output: $D_t = Y_t^{\mu}$, with $\mu$>0. Combining the two equations and taking logarithms, the following yields:

$$\ln P_t = \ln\left[\frac{1}{r_t + \theta - g}\right] + \ln Y_t^{\mu}. \tag{2.1.2}$$

Taking the derivate of this expression with respect to interest rate, we get:

$$\frac{\partial \ln P_t}{\partial r_t} = -\frac{1}{r_t + \theta - g} \tag{2.1.3}$$

Thus, an expansionary monetary policy which reduces the interest rate will increase asset prices.

The asset price misalignments, describing persistent and erratic deviations of asset prices from their fundamental values, which appeared in advanced economies before the eruption of global financial crisis have been related to the stance of monetary policy. In particular, the excessive easing of monetary policy which has brought policy rates at the zero bound as a way of stimulating the real economy has encouraged individuals and firms to borrow heavily, pushing up asset prices, thus creating a bubble-like behavior (Lane, 2016). The events which followed the global financial meltdown have indicated to governments and central bankers that the pursuit of financial stability should complement the pursuit of macroeconomic goals such as real growth and stable inflation rate (Yellen, 2014). However, one of the debatable issues in the practice of central banking currently is whether central banks should design monetary policy with respect to financial stability at the cost of other policy goals. At the theoretical level, the research agenda has addressed the link between monetary policy and financial stability in the context of the established macroeconomic model with financial frictions (Woodford, 2012; Brunnermeier and Sannikov, 2014). This link is very crucial for the stability and the functioning of financial system as an efficient allocation mechanism of savings to investments which promote economic growth and prosperity.

Several empirical studies have examined the monetary policy effects on individual asset returns, and the influence of macroeconomic and risk factors on stock-bond correlations. Given that asset price co-movements are of great importance to policy makers and portfolio investors, we add to literature on the determinants of asset price co-movements by examining the role of monetary policy developments in the United Kingdom (UK) in driving changes in dynamic correlations between stock and house returns, given the contribution of other macroeconomic variables. The importance of this topic stems from the fact that, firstly stock and house markets play an important role in the transmission mechanism of monetary policy, and secondly stocks and houses constitute large components of the total wealth of UK households, thus affecting consumption and investment decisions. According to the Office of National Statistics, at the end of 2015 for the household and non-profit institutions serving households sector dwellings were the most valuable non-financial asset at £5.5 trillion, accounting for 51% of its net worth, and the value of equity and investment fund shares and units was at £791 billion, accounting for 17% of this sector's net financial worth.

Initially, we use the dynamic conditional correlation (*DCC*) model proposed by Engle (2002) to estimate the covariance of stock-house returns. It is a multivariate generalized autoregressive conditional heteroskedasticity (GARCH) model in which correlations are dynamic according to an autoregressive specification. This parametric method is used extensively in the empirical literature to filter dynamic covariance between asset returns as it does not suffer from the computational difficulties of other multivariate GARCH models.

In turn, we use a least squares with breaks method developed by Bai-Perron (2003) to filter the variability of dynamic conditional covariance between the stock market return and house market return. The variability constitutes the cyclical component of dynamic covariance which reflects periods of positive and negative deviations from trend. If deviations are large, boom and bust outcomes are present and imply risk. Managing risk means having information about the whole distribution of these possible outcomes. To get information about these outcomes, we relate the variability of dynamic conditional covariance with developments in monetary policy. The modern portfolio theory together with the traditional asset pricing model offer a reasoning for this link, as we have previously discussed. Particularly, we examine whether predetermined monetary policy measures contain information for forecasting changes in the dynamic covariance of stock-house returns across the entire conditional distribution of the response variable. To this end, we use a quantile approach that is more informative to the research question than least squares which looks only at the conditional mean.[1] Modeling the stylized facts of dynamic co-movements of stock-house returns, the present study will provide policy makers with valuable information about monetary policy developments in characterizing these facts, and benefit investors in diversifying their portfolio strategies.

The rest of the chapter is organized as follows. Section 2.2 overviews the empirical literature. Section 2.3 presents a theoretical model which frames the link between the covariance of two asset returns and other variables. Section 2.4 presents the DCC method and the quantile approach which are used in the empirical analysis. Section 2.5 discusses the univariate properties of the data and the empirical results. Section 2.6 concludes.

## 2.2. Literature Review

There are two main strands of the literature which are associated with the present study. The first one relates individual asset market returns with the macroeconomic environment in general and monetary policy developments specifically, and the second one relates the covariance of asset returns with predetermined macroeconomic variables and risk factors which are derived from alternative asset pricing models. The main conclusions which resulted from the relevant literature is that firstly the stock and house markets constitute mechanisms though which monetary policy decisions affect the macroeconomic environment, and thus these facts should be taken into consideration when policy decisions are designed, and secondly the analysis of asset co-movements is mainly focused on equity and bond returns which are related to economic and other factors. This discussion reveals the gap in the literature referring to the co-movements of stock and house returns and the role of monetary

[1] The quantile methodology is discussed in Koenker and Bassett (1978), Koenker and Hallock (2001), Koenker and Xiao (2004), Koenker and Xiao (2005), Koenker (2005).

policy developments in the UK in driving changes to the dynamic correlations of these returns. Given that asset price co-movements are of great importance to policy makers and portfolio investors, this chapter aspires to fill this gap.

### 2.2.1. Monetary policy and asset prices

This part of the literature examines the link between monetary policy and asset markets using a plethora of econometric methods, including a simple regression analysis, an event-study analysis, and a multivariate time series analysis. We will begin the overview the first strand of the literature with a survey paper which provides information from two centuries. Bordo and Wheelock (2004) examined whether US stock market booms in the 19th and 20th centuries reflected a macroeconomic environment with real output growth and increases in productivity, inflation, expansionary monetary policies, or these events defy any explanation because of irrational exuberance. The analysis of two hundred years of data indicated that these events were mainly associated with real output growth and increases in productivity, thus suggesting the asset booms were related to fundamentals. Furthermore, expansionary monetary policies reflected in money and credit growth above average played a significant and positive role into the existence of the stock market booms. This finding reflects a passive accommodation of these episodes either by the banking system through the creation of additional loans, or by expansion of the monetary base especially in periods of gold inflows. It is worth mentioning that asset price booms occurred with inflation, deflation, or price stability. Bredin et al. (2007) examined how expected and unexpected monetary policy shocks in the UK from January 1993 to March 2004 affect the aggregate and sector stock market returns using an event study analysis. The expected change in the base rate is the difference between its actual change and its unexpected change which is proxied by the daily change in the three-month sterling LIBOR futures contract. At the aggregate and sectoral levels, the results indicated that a surprise fall in the base rate increased stock returns. On the other hand, the expected changes in the policy rate did not have any statistically significant effect on stock market returns. Finally, they documented evidence that unanticipated changes in the policy rate affect current excess return by changing expectations about future excess returns. Ioannidis and Kontonikas (2008) investigated the relationship between changes in monetary policy and stock market returns in thirteen OECD members from January 1972 to July 2002. Their findings suggested that in a large majority of the examined countries, when there were periods of strict monetary policy, a decrease in the value of stock market was observed then. This observation is consistent with the present value model. They also showed that not only current stock returns were influenced by monetary policy shocks, but also future ones. Thus, there was a negative relationship between strict monetary policy and future stock returns. As a result, following these observations, the authors suggested that monetary authorities are responsible for a suitable monetary policy rule, following changes in the stock market returns. Bohl et al. (2008) investigated how European stock market returns were instantly affected by sudden monetary policy changes, applied by the European Central Bank (ECB). For this purpose, they included the endogeneity between stock market returns and interest rate changes, and also the fact that the stock exchanges are well aware of the monetary policy of the ECB. They used daily data from January 1999 to February 2007 for the four biggest European stock markets (Germany, France, Spain and Italy). The Euro Stoxx 50 was used as a measure for the combined European market, and the one-month EURIBOR. The main empirical results suggested that there was a strong negative response from the European markets, when the ECB introduced a monetary policy shock and this response was spread evenly across the examined countries. Thus, the effects of a monetary policy change are immediate. In addition, their findings implied that monetary policy decisions were predictable, indicating that the ECB successfully communicated its policy decisions to

European stock markets. Konrad (2009) used a GARCH in mean model to examine the effects of monetary policy shocks by the Federal Reserve and the Bundesbank/ECB on the volatility of equity and bond returns in Germany over the period 1989-2007. He found that the volatility of equity return was attributable to monetary policy shocks by the Federal Reserve, while the volatility of bond return was explained by the surprises of ECB monetary policy. Furthermore, the surprise monetary policy effects on stock market volatility where larger during bearish than bullish phases. Joyce et al. (2010) investigated asset market reactions to news about quantitative easing (QE) purchases by the BoE from February 2009 to February 2010 using an event-study analysis. Their main results showed that the QE purchases had a strong impact on gilt yields through a portfolio-balance effect. In addition, large announcement effects were noticed for corporate yields, but no announcement effects were observed for the stock market. Gagnon et al. (2010) examined the implementation of long-term asset purchases (LSAPs) by the Federal Reserve since September 2008, and focused on how these purchases had an impact on the economy. The basic channel through which the LSAPs will affect the economy is the portfolio-balance effect. In particular, when the Federal Reserve buys long-term securities in exchange of high-powered money, private investors will be willing to adjust their portfolios when asset prices increase and hence asset yield fall. In other words, the reduction in asset supplies in private portfolios will reduce asset risk and consequently asset yields. The empirical analysis used an event-study where changes in interest rates in different public and private debt instruments were examined around the announcements of the LSAPs. Their results suggested that the LSAPs reduced significantly the long-term interest rates for a range of bonds, thus empowering the economic activity. Not only the mortgage market was benefited, but also the Treasury securities market, the corporate bonds market and the interest-rate swaps market. The latest result was important, as it was proved that even the zero bound was reached, the measure of asset purchases had a strong economic effect. Bauer and Neely (2012) examined how significant the signaling and portfolio balance channels were for the reduction of the international bond yields, following the implementation of Federal Reserve unconventional monetary policy. They used daily data for asset yields between January 1995 and September 2013, and their analysis was based on dynamic term structure models. Their empirical work suggested that both channels affected a lot the yields in most of the examined countries. In the U.S. and Canada, the signaling channel seemed to be very strong, while in Germany and Australia, the signaling channel had a weaker effect. On the other hand, in Japan, the effects from the signaling effect seemed to be very weak. Regarding the portfolio balance channel, strong effects were observed in Australia and Germany, weaker ones in the U.S. and Canada and quite modest in Japan. Gregoriou et al. (2012) examine the link between anticipated and unanticipated interest rate decisions by the BoE Monetary Policy Committee (MPC) on aggregate and sectoral asset returns using daily data from June 1999 to March 2009. The change in stock returns during the day of the meeting of MPC is regressed on the unanticipated change in the interest rate, which is proxied by the three-month sterling LIBOR for future contracts, and the anticipated change in the interest rate which is proxied by the difference between the actual change in the interest rate and the unanticipated rate. The empirical analysis indicated that the link between monetary policy actions and daily changes in stock returns was negative before the credit crisis, and turned positive during the credit crisis, thus indicating the inability of policy makers to reverse the fall in asset prices during the crisis period with interest rate reductions.

Rosa (2012) investigated how the LSAPs of the Federal Reserve had an effect on U.S. asset prices. After constructing the surprise part of the LSAPs announcements using information from Financial Times, then he used these announcements into a regression model. The results indicated that Federal Reserve asset purchases had a strong effect on U.S. asset prices. Thus, monetary policy and asset prices were closely connected, especially by a policy transmission mechanism. Goodhart and Hofmann (2008) investigated the link between house prices, monetary variables, real output, and inflation for seventeen industrialized countries from 1970 to 2006 using a panel vector autoregressive (VAR) model. The empirical results suggested the existence of strong multidirectional relationships between the examined variables. Moreover, these links seemed to be stronger during the last twenty years of their sample. Finally, the analysis indicated that shocks to monetary variables were stronger during periods when there was a booming in house prices. Elbourne (2008) examined how the housing market contributed to the monetary transmission mechanism in the UK over the period 1987 to 2003 using a structural VAR (SVAR) model with eight variables. In order to identify the short-run and medium-run effects of various shocks, he imposed restrictions on the contemporaneous matrix of the model based on macroeconomic theory. The results from the impulse response analysis indicated that housing price shocks, by affecting consumption, price level and interest rates, had indeed an influence into the transmission of monetary policy, but not as big as other authors have said. Iacoviello and Minetti (2008) examined the effectiveness of the credit channel of monetary transmission mechanism by focusing on four European housing markets (UK, Norway, Germany and Finland) using a VAR approach. They separated the bank-lending channel from the balance-sheet channel. The results for UK and Finland showed the existence of a bank-lending channel, for Germany the existence of a balance-sheet channel, whereas for Norway there was no strong evidence for any credit channel. In addition, the empirical evidence showed that residual heterogeneity characterized the housing markets of the four countries something which affected the transmission of monetary policy. Bjornland and Jacobsen (2010) examined whether house prices contributed to the monetary policy transmission mechanism in UK, Norway and Sweden using quarterly data from 1983 to 2006, and SVAR analysis. The selection of the variables was based on the structure of a New-Keynesian small open economy model, where domestic and foreign interest rates, inflation, real output, real house price, and real exchange rate were included. The empirical results showed that a monetary policy shock significantly reduced house prices, and this effect enhanced the negative effect of this shock on real output which observed in models without a housing market. As a result, housing market was a significant factor for the monetary transmission mechanism. In addition, house price shocks affected interest rates, but the dynamic pattern of this effect was not the same for all countries. McDonald and Stokes (2013) examined the origins of the bubble in the housing market in the United States using Granger causality and VAR analysis. The dataset consisted of the federal funds rate, and S&P/Case-Shiller Home Price Indices for ten and twenty cities respectively. The main results showed that during the period 2000-2010/8 there was Granger causality between the federal funds rate and the housing price. In addition, the results supported the view that the low interest rate policy during the period 2001-2004 was the main driver of the subsequent housing bubble. Tsai (2015) searched for the existence of any correlations between monetary policies and housing prices in the UK national and five regional markets using quarterly data from 1986 to 2011. This study used money supply and the short-term interest rate to represent the monetary policy variables. The co-integration approach indicated there was a long-run equilibrium relationship between the two monetary policy variables and house prices. Furthermore, the error correction model indicated that monetary policy affected the dynamics of housing market. The study claims that the strong correlation between the house price and monetary policy should be taken into consideration by the BoE so as to examine the effects of the implementation of any monetary policy into the housing market.

### 2.2.2. Co-movements in asset prices

The second strand of the empirical literature which is related to the present study has focused on asset return co-movements using time series analysis and multivariate GARCH approaches.

Shiller and Beltratti (1992) used a simple version of the present value model and a VAR methodology to calculate the theoretical correlations implied by the model between stock and bond markets. Exploiting annual data from the U.S. and the UK from 1871 to 1989, and from 1918 to 1989 respectively, they observed a negative correlation between stock prices and long-term interest rates. Furthermore, the results indicated a positive correlation between the actual excess returns in stock and bond markets. These results are not in line with the estimates of the present value model obtained from the VAR analysis. Longin and Solnik (1995) used a GARCH model with a constant conditional correlation to derive the dynamic pattern of correlations between excess returns for seven countries employing monthly data from 1960 to 1990. The results showed that the international correlation between the seven markets increased during the sample period, and this increase was associated with a rise in individual market volatility. In addition, the conditional correlation increased with an increase in the U.S. short-term interest rate and declined with the U.S. dividend yield. Kroner and Ng (1998) applied a general dynamic covariance model, which included the four mostly used multivariate GARCH models, in order to gather together their asymmetric applications, to extinguish any misspecifications these models have and finally to examine for any correlations between large and small firm returns. Their empirical findings suggested that large firm returns had an effect on the volatility of the small firm returns, but the opposite flow did not exist. In addition, when there was bad news about large firms, this led to volatility in all firms and also, the conditional covariance between small and large ones seemed to be higher than when there was good news. The paper concluded that what matters for asset pricing is the right choice of the correct multivariate volatility model. Hunter and Simon (2005) analyzed the conditional correlations between U.S. and U.K, Germany, and Japan bond returns using a bivatiate GARCH model with a constant correlation, and weekly data from January 1992 to September 2002. The results indicated that correlations between international bond returns were time-varying and affected by differences in business cycles conditions which were reflected on the yield curve slope differentials. Yang (2005) studied how the Japanese stock market was correlated with the stock markets of the Asian Four Tigers using daily data from 1990 to 2003, and the DCC approach. The empirical results indicated contagious effects across stock markets. Furthermore, he observed an increase in dynamic correlations when there were high market volatilities, which was discouraging for those with international portfolio diversification. Baele et al. (2010) examined how much of the time-varying co-movements between equity and bond markets returns is attributable to a number of economic and other factors. Using a dynamic factor model, they found that macroeconomic variables such inflation, output gap, and interest rates play a minor role in describing the correlations between equity and bond markets returns. On the other hand, liquidity factors explain a larger part of dynamic correlations. Gomes and Taamouri (2016) used an affine general equilibrium model to frame the relationship between the covariance of two asset returns and a number of latent state variables. Then, they employed the DDC model to filter the covariance between stock and bond markets in the Euro Area, and regression analysis to predict covariances with global and domestic factors which were extracted from principal component analysis from ten indices related to economy activity. The results showed that these factors explained half of the variation of covariances between stock returns, and about one third of variation of covariances between bond returns.

## 2.3. Theoretical Issues

In this section, we present an affine general equilibrium model developed by Gomes and Taamouti (GT, 2014 and 2016) which establishes a linear relationship between the covariance of asset returns and a vector of latent state variables. This linear relationship frames the empirical model used in the present study. We simplify their model by examining one economy with two assets *i* and *j*, and returns $r_{i,t+1}$ and $r_{j,t+1}$ respectively. The model assumes that in the endowment economy the representative household has a preference over a sequence of consumption paths $\{c_t, c_{t+1}, ..\}$, and these so-called Epstein-Zin-Weil preferences are described by the following recursive utility form:

$$U_t = \left[ (1-\beta) c_t^{\gamma} + \beta \left( E_t U_{t+1}^{\alpha} \right)^{\frac{\gamma}{\alpha}} \right]^{\frac{1}{\gamma}} \qquad (2.3.1)$$

where the parameter $0<\beta<1$ denotes the household's discount rate, the parameter $\gamma<1$ determines the elasticity of intertemporal substitution $\zeta = 1/(1-\gamma)$, the parameter $\alpha<1$ denotes the risk aversion, and $E_t U_{t+1}$ is the expected value of the future utility, with expectations formed at time *t*. GT assume that the joint dynamics of the growth rate of consumption, $g_{t+1}$, and a *n*-dimensional vector of latent state variables $x_{t+1}$ of the economy has the Laplace transformation:

$$E_t\left[ e^{s_c g_{t+1} + v'_x x_{t+1}} \right] = e^{F_0(s_c, v_x) + x' F_x(s_c, v_x)} \qquad (2.3.2)$$

where $F_0$ is a scalar function which assigns a scalar to $s_c$ and $v_x$, and $F_1$ is a vector function which assigns a vector *x* to $s_c$ and $v_x$. GT model the behavior of $r_{i,t+1}$ and $r_{j,t+1}$ using Campbell-Shiller approximation: $r_{k,t+1} = \varphi_{k,0} + \varphi_{k,1} z_{t+1} - z_t + g_{t+1}$, $k=i,j$, where $z_t$ is the wealth-consumption ratio which follows the process: $z_t = A + B' x_t$ (see Eraker, 2006). Using equations (2.3.1) and (2.3.2), GT derive the joint conditional cumulant-generating function of $r_{i,t+1}$ and $r_{j,t+1}$, $\Gamma(s_i,s_j)$, which is an affine function of the $n$-dimensional vector of $x_{t+1}$ variables, with the following form,

$$\Gamma(s_i,s_j) = \log\left\{E_t\left(e^{s_i r_{i,t+1}+s_j r_{j,t+1}}\right)\right\} = \overline{F}_0(s_i,s_j) + x'_t\,\overline{F}_x(s_i,s_j) \qquad (2.3.3)$$

where $\overline{F}_0(s_i,s_j) = F_0(s_i,s_j) + s_i(\varphi_{i,0} + A\varphi_{i,1} - A) + s_j(\varphi_{j,0} + A\varphi_{j,1} - A)$, and

$\overline{F}_x(s_i,s_j) = F_x(s_i,s_j) - B's_i + B's_j$, with $F_0(s_i,s_j) \equiv F_0(s_i + s_j, B'\varphi_{i,1}s_i + B'\varphi_{j,1}s_j)$, and

$F_x(s_i,s_j) \equiv F_x(s_i + s_j, B'\iota\varphi_{i,1}s_i + B'\iota\varphi_{j,1}s_j)$, where $\iota$ is an $n$-dimensional vector of ones. Using equations (2.3.1), (2.3.2) and (2.3.3), GT show that the covariance of the two asset returns $r_{i,t+1}$ and $r_{j,t+1}$ is a linear function of the state variables $x_{t+1}$,

$$E_t(r_{i,t+1}r_{j,t+1}) = \delta_0 + x'_{t+1}\,\delta_x \qquad (2.3.4)$$

where $\delta_0 = \dfrac{\partial^2\left[\overline{F}_0(s_i,s_j)\right]}{\partial s_i \partial s_j}$, $\delta_x = \dfrac{\partial^2\left[\overline{F}_x(s_i,s_j)\right]}{\partial s_i \partial s_j}$.

While several studies associate the latent state variables with predetermined variables such as inflation, real output and divided-yield ratio to explain the covariance of asset returns, GT use instead a principal component analysis to extract a number of factors from ten indicators of economic activity and employ these factors to predict the covariance of asset returns.

An implication of the theoretical model for the empirical analysis of this chapter is that the covariance of the two asset returns is a linear function of state variables. In this chapter, we associate the latent state variables with constructed-on-purpose moving average monetary variables, such as monetary aggregates and the interest rate, to predict the whole distribution of the conditional covariance of asset returns, given the presence of other macroeconomic variables. To this end, we use two empirical approaches. Firstly, we employ the DCC method to filter the dynamic covariance between the stock and house returns, and secondly, we use the quantile technique to examine the predictive content of monetary policy variables on the entire conditional distribution of dynamic covariance fluctuations.

## 2.4. Empirical Methodology

In this chapter, we use two empirical approaches to address two issues. Firstly, we employ the DCC method to filter the dynamic covariance between the stock and house returns, and secondly, we use the quantile technique to examine the predictive content of monetary policy variables on the entire conditional distribution of fluctuations of dynamic covariance. In turn, we will explain both methodologies.

### 2.4.1. The dynamic conditional correlation approach

Let us assume that $r_{i,t}$ and $r_{j,t}$ denote, respectively, the stock market and house market returns, defined as $r_{i,t} = \ln(sp_t / sp_{t-1})$, $r_{j,t} = \ln(hp_t / hp_{t-1})$, and $r_t = (r_{i,t}, r_{j,t})'$ be a 2*x*1 vector of financial returns which follows a stationary process of the form,

$$r_t | \mathrm{H}_{t-1} \sim N(0, \Sigma_t) \tag{2.4.1}$$

where $\mathrm{H}_{t\text{-}1}$ is the information set at time *t*-1 used to explain $r_t$ , and $\Sigma_t = \begin{bmatrix} \sigma_{i,t}^2 & \sigma_{ij,t} \\ \sigma_{ji,t} & \sigma_{j,t}^2 \end{bmatrix}$ is the time-varying covariance matrix, with $\sigma_{ij,t} = \sigma_{ji,t}$.

Let the conditional variances of the two financial returns, $\sigma_{i,t}^2$ and $\sigma_{j,t}^2$, be described by the univariate GARCH model:

$$r_{k,t} = \varepsilon_{k,t}, \qquad \varepsilon_{k,t} \sim N(0, \sigma_{k,t}^2) \tag{2.4.2}$$

$$\sigma_{k,t}^2 = \psi_0 + \psi_1 \sigma_{k,t\text{-}1}^2 + \psi_2 \varepsilon_{k,t-1}^2 \tag{2.4.3}$$

where $r_{k,t}$ is the financial return of the *k*-th market (*k*=*i*,*j*), $\sigma_{k,t}^2$ and $\varepsilon_{k,t}^2$ are respectively the conditional and unconditional variances of *k*-th return. The first expression describes the mean equation which assumes that the financial return is a random walk with zero mean (that is, the intercept of the model is zero), and the second expression describes the conditional variance of the financial return.

Let $D_t$ be a diagonal matrix with the two conditional standard deviations (volatilities) on the diagonal, $D_t = \begin{bmatrix} \sigma_{i,t} & 0 \\ 0 & \sigma_{j,t} \end{bmatrix}$, and $R_t$ be the time-varying correlation matrix of the two financial returns, $R_t = \begin{bmatrix} 1 & \rho_{ij,t} \\ \rho_{ji,t} & 1 \end{bmatrix}$. Then, the time-varying covariance matrix $\Sigma_t$ is equal to $D_t R_t D_t$. Expanding this formula, we get:

$$\Sigma_t = \begin{bmatrix} \sigma_{i,t} & 0 \\ 0 & \sigma_{j,t} \end{bmatrix} \begin{bmatrix} 1 & \rho_{ij,t} \\ \rho_{ji,t} & 1 \end{bmatrix} \begin{bmatrix} \sigma_{i,t} & 0 \\ 0 & \sigma_{j,t} \end{bmatrix} = \begin{bmatrix} \sigma_{i,t} & 0 \\ 0 & \sigma_{j,t} \end{bmatrix} \begin{bmatrix} \sigma_{i,t} & \rho_{ij,t}\sigma_{j,t} \\ \rho_{ji,t}\sigma_{i,t} & \sigma_{j,t} \end{bmatrix}$$

$$= \begin{bmatrix} \sigma_{i,t}^2 & \sigma_{i,t}\rho_{ij,t}\sigma_{j,t} \\ \sigma_{j,t}\rho_{ji,t}\sigma_{i,t} & \sigma_{j,t}^2 \end{bmatrix} = \begin{bmatrix} \sigma_{i,t}^2 & \sigma_{i,t}\dfrac{\sigma_{ij,t}}{\sigma_{i,t}\sigma_{j,t}}\sigma_{j,t} \\ \sigma_{j,t}\dfrac{\sigma_{ji,t}}{\sigma_{j,t}\sigma_{i,t}}\sigma_{i,t} & \sigma_{j,t}^2 \end{bmatrix} = \begin{bmatrix} \sigma_{i,t}^2 & \sigma_{ij,t} \\ \sigma_{ji,t} & \sigma_{j,t}^2 \end{bmatrix}$$

We construct the standardized residuals series $\omega_{k,t}$ as a $N(0,1)$ process by scaling the residuals $\varepsilon_{k,t}$ of model (2.4.2) by their conditional volatility, that is $\omega_{k,t} = \varepsilon_{k,t} / \sigma_{k,t}$. The conditional correlation of standardized residuals is the ratio of conditional covariance to the product of two conditional variances: $\rho_t = cov(\omega_{i,t}\omega_{j,t}) / var(\omega_{i,t}) var(\omega_{j,t})$. Each term in the denominator is equal to one, and thus the time-varying conditional correlations are also the conditional covariance between the standardized residuals: $\rho_t = cov(\omega_{i,t}\omega_{j,t})$. The DCC model for two financial returns with a GARCH(1,1) specification is the following:

$$\rho_t = (1-\vartheta_1-\vartheta_2)\bar{\rho}_t + \vartheta_1(\omega_{i,t-1}\omega_{j,t-1}) + \vartheta_2\rho_{t-1} \qquad (2.4.4)$$

where $\bar{\rho}_t$ describes the unconditional correlation between the standardized residuals $\omega_{i,t}$ and $\omega_{j,t}$, and $\vartheta_1$ and $\vartheta_2$ are non-negative scalars with $\vartheta_1+\vartheta_2\langle 1$, if the model is a mean reverting process. A two-step approach is followed to estimate the DCC model (2.4.4). First, a univariate GARCH model is fitted to each financial return, and then the parameters estimates of GARCH models are used as inputs to obtain the parameters of the DCC model.

The stationarity of the constructed series $\rho_t$ has been examined using a breakpoint Dickey-Fuller (*DF*) unit root test and the Bai-Perron (2003) multiple break test. In the DF test, the breaking selection is to choose the model which minimizes the DF *t*-statistic. The innovation outlier model is specified as follows:[2]

$$\rho_t = a_0 + a_1 t + \alpha_2\rho_{t-1} + a_3 DU_t(T_b) + a_4 D_t(T_b) + \alpha_5 DT_t(T_b) + \varepsilon_t \qquad (2.4.5)$$

where $T_b$ is the uncertain break date, $DU_t(T_b)$ is an intercept break dummy variable that takes 0 for all dates before the break and 1 afterwards, $D_t(T_b)$ is one-time break dummy variable that takes 1 on the break date and 0 elsewhere, and $DT_t(T_b)$ is a trend break dummy variable which takes 0 for all dates before the break and values 1, 2, 3, etc. thereafter. The DF test statistic is the *t*-statistic for testing the null hypothesis that $\rho_t$ has a unit root, $a_2 = 1$, against the alternative hypothesis that $\rho_t$ is a stationary process, $|a_2| < 1$, with a break at $T_b$.

The Bai-Perron regression, in our case, has the form:

$$\rho_t = X_t'\delta + e_t \qquad (2.4.6)$$

[2] For a survey of the literature, see Perron (2006).

where the vector $X_t$ contains a constant and a time trend whose coefficients $\delta$ are subject to regime-change, and $e_t$ is the error term. The time series $\rho_t$ is subject to $n$ structural breaks permitting $n$+1 partitions of the series. They test the hypothesis of no structural breaks, $\delta_0=\delta_1=...=\delta_{k+1}$, against the alternative hypothesis of a number of $n$=$k$ breaks using the $F$-test:

$$F(\hat{\delta}) = \frac{1}{T}\left(\frac{T-(k+1)q-p}{kq}\right)(R\hat{\delta})'\left(RV(\hat{\delta})R'\right)^{-1}R\hat{\delta} \qquad (2.4.7)$$

where $\hat{\delta}$ is the $k$-break estimator of $\delta$, $q$ is the maximum number of observations each of the $n$+1 partitions of the series should have, $p$ is the number of regressors, $(R\hat{\delta})' = (\delta_0' - \delta_1', \ldots, \delta_k' - \delta_{k+1}')$, and $V(\hat{\delta})$ is an estimate of the variance-covariance estimator of $\hat{\delta}$. An additional $F$-statistic is proposed to test the null hypothesis of $k$ breaks against the alternative hypothesis of $k$+1 breaks using a sequential procedure over the $n$+1 segments of the series. The null hypothesis is not accepted if the global minimum of the sum of squared residuals in all segments of the series is smaller than sum of squared residuals of the model with $k$ breaks.

### 2.4.2. The quantile approach

The positive and negative changes in the joint fluctuations of stock and house returns can be associated with extreme quantiles of their conditional distribution. The quantile methodology can be used to examine whether these changes in the entire conditional distribution of stock and house returns are related to predetermined monetary policy variables, and this is the reason why we have selected this approach as opposed to least squares method which looks only at the mean effects.[3]

Let us assume that the variability of the dynamic correlation of stock-house returns is determined by the process: $\widetilde{\rho}_t = \lambda_1 \widetilde{\rho}_{t\text{-}1} + \lambda_2 x_{t-1} + \xi_t$, where $x_{t-1}$ is a predetermined variable which predicts the response variable, and $\xi_t$ is the error term. If $Q_{\widetilde{\rho}_t}\left(\tau \middle| \widetilde{\rho}_{t\text{-}1}, x_{t-1}\right)$ denotes the *τ-th* conditional quantile of $\widetilde{\rho}_t$ conditional on $\widetilde{\rho}_{t\text{-}1}, x_{t-1}$, and $Q_{\xi}(\tau)$ denotes the *τ-th* quantile of $\xi_t$, then $Q_{\widetilde{\rho}_t}\left(\tau \middle| \widetilde{\rho}_{t\text{-}1}, x_{t-1}\right) = Q_{\xi}(\tau) + \lambda_1 \widetilde{\rho}_{t\text{-}1} + \lambda_2 x_{t-1}$.

Let $\vartheta_0(\tau) = Q_{\xi}(\tau)$, $\vartheta_1(\tau) = \lambda_1$, $\vartheta_2(\tau) = \lambda_2$, and define $\vartheta(\tau) = \left(\vartheta_0(\tau), \vartheta_1(\tau), \vartheta_2(\tau)\right)'$,

then the quantile autoregressive model which we will use to answer the research question is given by:

$$Q_{\widetilde{\rho}_t}\left(\tau \middle| \widetilde{\rho}_{t\text{-}1}, x_{t\text{-}1}\right) = \left(1, \widetilde{\rho}_{t\text{-}1}, x_{t-1}\right)' \begin{pmatrix} \vartheta_0(\tau) \\ \vartheta_1(\tau) \\ \vartheta_2(\tau) \end{pmatrix} \tag{2.4.8}$$

The *τ-th* conditional quantile is defined as the value $Q_{\widetilde{\rho}_t}\left(\tau \middle| z_{t-1}\right)$, where $z_{t-1} = \left(\widetilde{\rho}_{t-1}, x_{t-1}\right)$, such that the probability that $\widetilde{\rho}_t$ conditional on $z_{t-1}$ will be less than this value is *τ*, and the probability that it will be more than this value is 1−*τ*. The linear relationship between the response variable $\widetilde{\rho}_t$ and the regressor $z_{t-1}$ at specified quantiles is estimated with the least absolute deviations (LAD) estimator which is a robust method as it does not imposes strong distributional assumption of the error term. The LAD estimator minimizes the sum of absolute residuals that gives asymmetric penalties 1-*τ* for over prediction and *τ* for under

[3] The quantile methodology is discussed in Koenker and Bassett (1978), Koenker and Hallock (2001), Koenker and Xiao (2004), Koenker (2005), Koenker and Xiao (2006).

prediction. The quantile regression estimator minimizes the following objective function (Baum, 2013):

$$Q_{\tilde{\rho}_t}\left(\vartheta(\tau)\right) = \sum_{t=1:\tilde{\rho}_t > z'_t \vartheta}^{T} \tau\left|\tilde{\rho}_t - z'_t \vartheta\right| + \sum_{t=1:\tilde{\rho}_t < z'_t \vartheta}^{T} (1-\tau)\left|\tilde{\rho}_t - z'_t \vartheta\right|$$

The null hypothesis that the autoregressive coefficient $\tilde{\rho}_{t-1}$ is equal to one against the alternative hypothesis that the coefficient is less than one can be tested using the covariate augmented DF (CADF) *t*-test (Hansen, 1995). The asymptotic distribution of this test depends on the squared correlation coefficient: $\delta^2 = \sigma^2_{\hat{\varpi}\hat{\xi}} \big/ \sigma^2_{\hat{\varpi}}\, \sigma^2_{\hat{\xi}}$, where $\hat{\xi}$ are the estimated residuals from the quantile equation (2.4.8) and $\hat{\varpi}$ are the estimated residuals from a restricted version of this equation which excludes the covariate *x*. Hansen has tabulated asymptotic critical values for the *t*-statistic ($t_n$) at different values of $\delta^2$ which can be used in the quantile regression at a fixed quantile $\tau$ ($t_n(\tau)$).

Initially, in the quantile equation (2.4.8) the vector $x_{t-1}$ includes only the monetary policy measures, which constitute the focus of our research question. In order to check the robustness of the obtained results, it is important to control for the impact that other predictors considered in the literature may have on the variability of dynamic correlations of stock-house returns. Thus, we have estimated a multivariate version of equation (2.4.8) which, except from the monetary variables, also includes other macroeconomic variables. However, the existence of possible correlation between these macroeconomic variables diminishes seriously their significance to act as additional predictors in the quantile regression. To this end, we use a Principal Component Analysis (PCA) to extract a lower dimensional space of linear combinations (the so-called principal components) between the additional control variables which are uncorrelated from each other and explain the bulk of the variance of the original data. The selected principal components will be utilized as additional indicators in the quantile regression.

Let a random vector $x$ consists of $j$ variables $k_1$, ...,$k_j$, that is $x' = (k_1,\ldots,k_j)$, and the variance-covariance matrix:

$$\Sigma = \begin{bmatrix} \sigma_1^2 & . & \sigma_{1j} \\ . & . & . \\ \sigma_{j1} & . & \sigma_j^2 \end{bmatrix}.$$

We consider the linear combinations between the $j$ variables: $x_1 = \varphi_{11}k_1 + \ldots + \varphi_{1j}k_j$, $x_2 = \varphi_{21}k_1 + \ldots + \varphi_{2j}k_j$, and so on until $x_j = \varphi_{j1}k_1 + \ldots + \varphi_{jj}k_j$. The PCA selects the coefficients $\varphi_{11}$, ..,$\varphi_{1j}$ that maximize the variance $var(x_1) = \varphi_1'\Sigma\varphi_1$ subject to the constraint $\varphi_1'\varphi_1 = 1$, the coefficients $\varphi_{21}$, ..,$\varphi_{2j}$ that maximize the variance $var(x_2) = \varphi_2'\Sigma\varphi_2$ subject to the constraint $\varphi_2'\varphi_2 = 1$, and so on until the coefficients $\varphi_{j1}$, ..,$\varphi_{jj}$ that maximize the variance $var(x_j) = \varphi_j'\Sigma\varphi_j$ subject to the constraint $\varphi_j'\varphi_j = 1$. The solution of this procedure produces the eigenvalues and eigenvectors of Σ.

## 2.5. Empirical Analysis

### 2.5.1. Data

In this chapter, we utilize monthly data from January 1983 to December 2014. The end of the sample has been dictated by the fact that there is no more recent data available for the Halifax price index. The use of high frequency monthly data ensures the presence of GARCH effects in stock and house returns. In order to create the two asset returns, we use the total share price index for all shares (SP) and the Halifax price index (HP).[4] In addition, we use three measures of monetary policy measures: the broad monetary aggregate M4 which refers to monetary financial institutions' sterling liabilities to private sector, the narrow monetary aggregate M0 which includes the average amount outstanding of total sterling notes and coin in circulation, excluding backing assets for commercial banknote issue in Scotland and Northern Ireland, and the official interest rate. The two monetary aggregates have been used as indicators of monetary policy and are the only aggregates for which monitoring ranges have been set by the BoE (Breedon and Fisher, 1994). These monitoring ranges have been established with the introduction of an inflation target and the use of the base interest rate as the primary monetary policy instrument to achieve inflation objective (OECD, 1998). We have also used the inside money (IM) which is constructed after subtracting M0 from M4. The monetary aggregates and the interest rate have been obtained from the BoE interactive Data Base, the share price from the Federal Reserve Economic Data (FRED) St. Louis, and the house price form the Halifax Building Society. The variables apart from the interest rate are firstly transformed to natural logarithms, and then to growth rates by taking the first difference of the logged levels. The change in the interest rate is computed as its first difference. In order to account for the delay in the pass-through of changes in the monetary policy measures on asset markets, we follow Rudebusch and Svensson (1999) and construct the monetary policy measures as lagged moving averages. The money measure is defined as the *k*-month moving average of monthly growth rates of money, $M(k)=1/k\left(\sum_{q=1}^{k}\Delta m_{t-q}\right)$, where $m_t$ denotes the logarithm of the level of a monetary aggregate and $\Delta m_t$ denotes its growth rate at time *t*. The interest rate measure is defined as the *k*-month moving average of monthly changes in the official rate, $\Delta R(k)=1/k\left(\sum_{q=1}^{k}\Delta R_{t-q}\right)$. We have constructed the three-

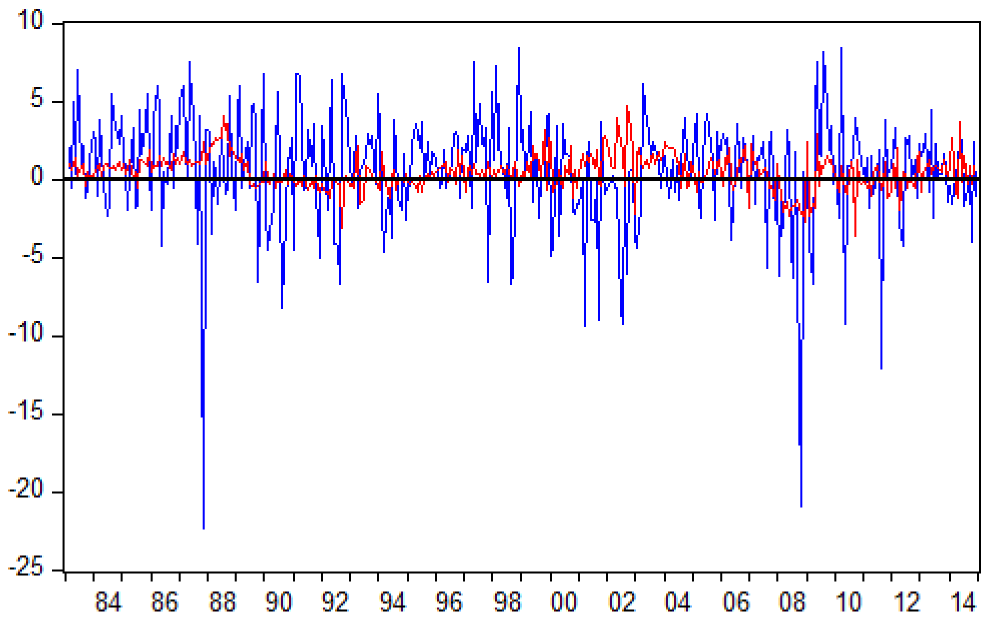
[4] Elbourne (2008) discusses the advantages of the HP index against other available house price indices.

month and six-month ($k$=3, 6) moving averages of monthly growth rates of M4, M0 and IM, and the three-month and six-month moving averages of monthly changes in the official rate.

### 2.5.2. Univariate analysis

The first two columns of Table 2.1 present the summary statistics of the original returns of stock and house markets, denoted as $r_{i,t}$ and $r_{j,t}$. The first return depicts the growth rate of stock prices and the second return depicts the growth rate of house prices. From Figure 2.1, we observe that the stock market return has been more volatile than the house market return.

Figure 2.1: The stock and house returns (original series)




Dickey and Fuller (1979), Philips and Perron (1998), and Kwiatkowski et al. (1992) unit root tests indicate that the stock market return and house market return are stationary. The first two test the hypothesis that the series contain a unit root, whereas the third one tests the hypothesis that the series are stationary. In Table 2.1 we report the Kwiatkowski et al.

(KPSS) test. The sample distribution of $r_{i,t}$ and $r_{j,t}$ is non-normal, as evidenced by high excess kurtosis and highly significant Jarque-Bera statistics.

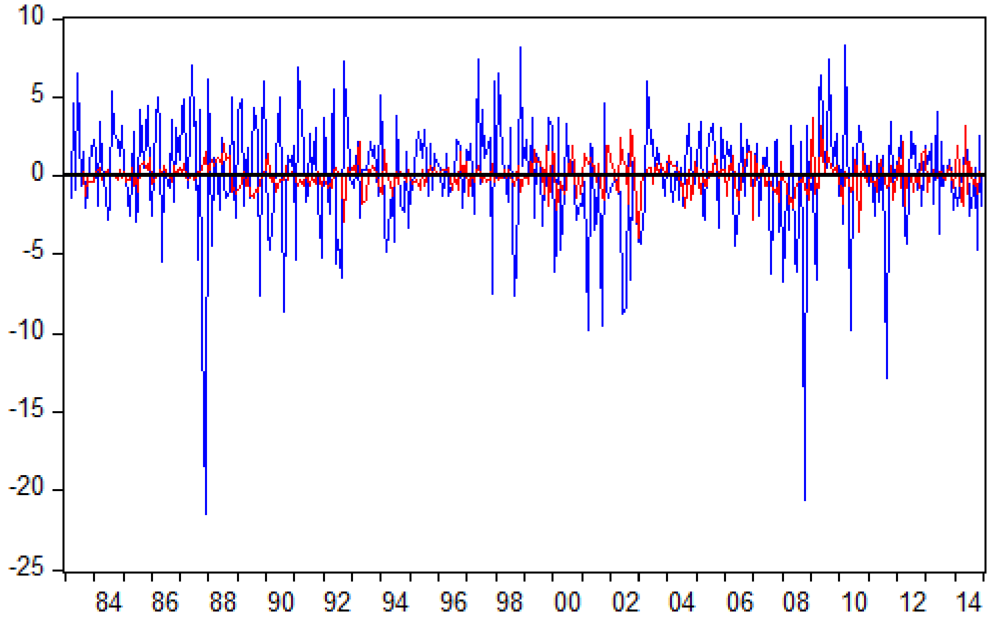
**Table 2.1: Summary statistics on stock and house market returns**

| | $r_i$ | $r_j$ | $e_i$ | $e_j$ |
|---|---|---|---|---|
| Mean | 0.543 | 0.484 | 0.000 | 0.000 |
| Median | 0.928 | 0.528 | 0.371 | 0.021 |
| Std. Dev. | 3.593 | 1.145 | 3.556 | 0.957 |
| Skewness | -1.598 | 0.058 | -1.419 | 0.014 |
| Ex. Kurtosis | 10.244 | 4.273 | 9.203 | 4.941 |
| Jarque-Bera | 1000.518*** | 26.054*** | 740.804*** | 59.322*** |
| KPSS | 0.370** | 0.201** | 0.354** | 0.084** |

**Notes:** The 5% asymptotic critical value of the LM statistic of KPSS test is equal to 0.463.***, ** indicate significance at 1% and 5% levels respectively.

To account for the presence of serial correlation in $r_{i,t}$ and $r_{j,t}$, we have whitened the two asset returns. In particular, we have modeled each return as a six-order autoregressive process and have used the Schwartz information criterion (SIC, Schwarz, 1978) to select the optimal lag structure. This criterion is defined as: $SIC = k\ln(n)/n - 2\ln(\hat{L})/n$, where $n$ is the sample size, $k$ is the number of regressors, and $\hat{L}$ is the value of the likelihood function which is maximized by the estimation of the model. We select the model with the lowest SIC. The optimal lag is one for the stock market return and five for the house market return. The selected autoregressive lag models have been tested for serial correlation over a number of lags using the Ljung-Box Q-statistic which is distributed as $\chi^2(q)$ with $q$ the degrees of freedom representing the number of lags being tested (Ljung and Box, 1978). The null hypotheis is that the correlations are zero and thus the data are independently distributed against the alternative hypothesis that the data exhibit serial correlation. The value of the Ljung-Box Q-statistic with 10 degrees of freedom for the stock return is equal to 8.634 with $p$-value=0.566 and for the house return is equal to 2.242 with $p$-value=0.994, and thus the null hypothesis of non serial correlation is not rejected. The estimated residuals from each model constitute the whitening series of financial returns, denoted as $e_{i,t}$ and $e_{j,t}$. The summary statistics of the two variables are presented in Table 2.1. The stock return has the biggest volatility, as evident from Figure 2.2.

Figure 2.2: The stock and house returns (whitening series)




The unit root test indicates that the two financial returns are stationary. The sample distributions of $e_{i,t}$ and $e_{j,t}$ are non-normal, as evidenced by high excess kurtosis and highly significant Jarque-Bera statistics. Thus, the bivariate GARCH analysis of the two asset returns will assume a $t$-student's error distribution.

### 2.5.3. GARCH-DCC analysis

We have fitted different univariate GARCH models to the financial returns $e_{i,t}$ and $e_{j,t}$, and the analysis has indicated that a GARCH(1,1) model fits adequately the data generation process of the two time series. We have estimated the GARCH-DCC model with maximum likelihood and have used *t*-studentst's distribution as the density function for standardized residuals $\omega_t$. Given the model: $e_{k,t} = c + \varepsilon_{k,t}$ and $\sigma^2_{k,t} = \psi_0 + \psi_1 \sigma^2_{k,t\text{-}1} + \psi_2 \varepsilon^2_{k,t-1}$, with $\varepsilon_{k,t} \sim N\left(0, \sigma^2_{k,t}\right)$ and $\omega_{k,t} = \varepsilon_{k,t} / \sigma_{k,t}$, the form of the conditional density function of the standardized residuals is given by:

$$f\left(\omega_{k,t} \middle| H_{t-1}\right) = \frac{\Lambda(v)}{\sqrt{\sigma^2_{k,t}}} \left[1 + \frac{\omega^2_{k,t}}{v-2}\right]^{-(v+1)/2}$$

with $\Lambda(v) = \dfrac{\Gamma\left[(v+1)/2\right]}{\sqrt{\pi (v-2)} \Gamma(v/2)}$. The degrees of freedom (DOF) parameter $v > 2$ controls the tail behavior. The *t*-student's distribution is normal as $v \to \infty$. A logarithmic version of the above formula is discussed in Eviews 9.5. The DOF estimated parameter *v* is small, and thus the distribution of the standardized residuals is far from normal. Table 2.2. reports the results from the GARCH and DCC modelling.

| **Table 2.2: GARCH and DCC modelling of stock and house returns** | | | | |
|---|---|---|---|---|
| | | **GARCH model** | | |
| | Stock market return | | House market return | |
| Regressors | Mean | Varinace | Mean | Variance |
| Constant | 0.276** (0.142) | 4.376*** (1.792) | 0.009 (0.04) | 0.081** (0.038) |
| $\sigma_{t-1}^{2}$ | | 0.399** (0.18) | | 0.634*** (0.092) |
| $\varepsilon_{t-1}^{2}$ | | 0.290*** (0.114) | | 0.326*** (0.105) |
| DOF parameter *v* | 4.479*** (1.126) | | 6.828*** (2.16) | |
| Serial correlation for standardized residuals | Q(30)=37.961 [*p*-value=0.151] | | Q(30)=41.920 [p-value=0.073] | |
| ARCH test for standardized residuals | F(1,375)=0.004 [*p*-value=0.951] | | F(1,375)=0.3767 [*p*-value=0.5397] | |
| Regressors | **DCC model** | | | |
| $\vartheta_1$ | 0.018* (0.01) | | | |
| $\vartheta_2$ | 0.974*** (0.02) | | | |
| DOF parameter *v* | 6.511*** (1.112) | | | |

**Notes:** Sample size: 384. The stability condition $\vartheta_0 + \vartheta_1 < 1$ is met. Bollerslev-Wooldridge robust standard errors and covariance for univariate fits were used. ***, **, * denote, respectively, significance at 1%, 5% and 10% levels.

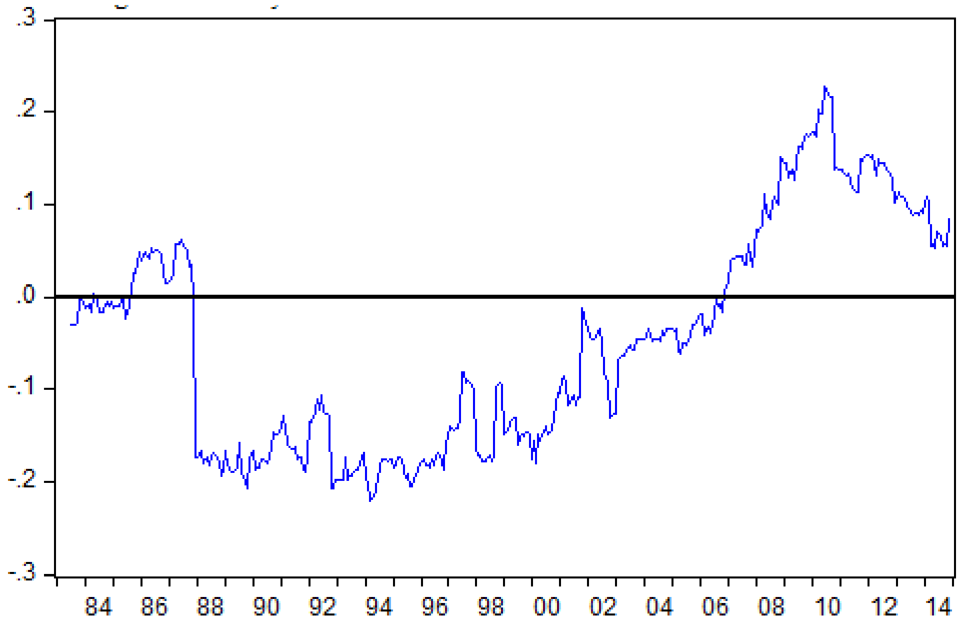
The *Q*-statistics which test the null hypothesis of serial correlation in the standardized and square standardized residuals, and the *F*-statistic which test the null hypothesis that the standardized residuals do not have first-order autoregressive conditional heteroskedasticity (ARCH) effect indicate that the GARCH models are well specified. After obtaining the conditional variances of the two financial returns, then we estimate the dynamic conditional correlations between $e_{i,t}$ and $e_{j,t}$, denoted as $\rho_t$.The results show that the stability condition $\vartheta_1+\vartheta_2<1$ is met, indicating that the *DCC* model is mean-reverting, which however exhibits a high degree of persistence in the conditional correlation, as the sum of the two coefficients is very close to unity. A breakpoint *DF* unit root test with an innovation outlier break type is used to test the hypothesis that $\rho_t$ is a stationary process. The model includes an intercept and a time trend. The breaking selection is to choose the model which minimizes the DF *t*-statistic. We have augmented the DF equation with 16 lagged values of the dependent variable and used the SIC to select the optimal lag length. The test has selected zero lags. Initially, we test hypothesis that the model exhibits a break in the intercept, and then we examine the hypothesis that the model exhibits a break in both the intercept and the time trend. The results are presented in the second panel of Table 2.3. The DF test statistic equals -5.4707 which is smaller than the 1% critical value, thus rejecting a unit root, with a break date being located at 1987M10, when the world economy experience the stock market crash. The trend break dummy variable is not statistically significant whereas the intercept break dummy variable is still statistically significant. Thus, the evidence from the modified DF unit root test reveals that the level of conditional correlation is a stationary process with a break date being located at 1987M10, and exhibits a mean-reversion with high persistence, as its autoregressive root is about 0.91.

| Table 2.3: The breakpoint *DF* unit root test for dynamic correlations | | |
|---|---|---|
| Regressors | Break in intercept | Break in intercept and trend |
| Constant | -0.001<br>(0.003) | 0.001<br>(0.005) |
| Trend | 0.001***<br>(2.2E-05) | 8.6E-05<br>(0.001) |
| $\rho_{t-1}$ | 0.909***<br>(0.017) | 0.909***<br>(0.017) |
| $DU(T_b)$ | -0.030***<br>(0.006) | -0.029***<br>(0.007) |
| $D(T_b)$ | 0.030*<br>(0.019) | 0.030*<br>(0.018) |
| $DT(T_b)$ | | 3.4E-05<br>(0.0001) |

**Notes:** Sample size: 384. Dependent variable is the dynamic conditional correlations between $e_{i,t}$ and $e_{j,t}$, denoted as $\rho_t$. Numbers are estimated coefficients. Standard errors (SEs) are reported in parentheses. ***, * indicate significance at 1% and 10% levels respectively. The regression with a break in the intercept is given by: $\rho_{ij,t} = \alpha_0 + a_1 t + a_2\rho_{ij,t-1} + a_3 DU_t(T_b) + a_4 D_t(T_b) + \varepsilon_{ij,t}$,and the tested hypothesis is that $\rho_{ij,t}$ has a unit root, $\alpha_2 = 1$, against the alternative hypothesis that $\rho_{ij,t}$ is a stationary process, $|\alpha_2| < 1$, with a break in intercept at the uncertain date $T_b$. The regression with a break in the intercept and the trend is given by: $\rho_{ij,t} = \alpha_0 + a_1 t + a_2\rho_{ij,t-1} + a_3 DU_t(T_b) + a_4 D_t(T_b) + \alpha_5 DT_t(T_b) + \varepsilon_{ij,t}$, and the ested hypothesis is that $\rho_{ij,t}$ has a unit root, $\alpha_2 = 1$, against the alternative hypothesis that $\rho_{ij,t}$ is a stationary process, $|\alpha_2| < 1$, with a break in intercept and the trend at the uncertain date $T_b$. $DU_t(T_b)$ is an intercept break dummy variable that takes 0 for all dates before the break and 1 afterwards, $D_t(T_b)$ is one-time break dummy variable that takes 1 on the break date and 0 elsewhere, and $DT_t(T_b)$ is a trend break dummy variable which takes 0 for all dates before the break and values 1, 2, 3, etc thereafter. The *DF* test statistic is the *t*-statistic for testing the null hypothesis of a unit root.

In Figure 2.3, which plots the dynamic conditional correlations between stock market and house market returns, it is evident that correlations are time-varying and behave rather heterogeneously. In particular, they were negative from the world stock market crash in October 1987 until the end of 2006, and then they were positive. Initially, a sharp decline occurred at November 1987 when the stock price index in the UK attained its lowest level as a result of the world stock market crash, indicating that stock and house returns moved opposite. Since the beginning of 2007, the correlations became positive, even during the global financial crisis 2007-2009, indicating that the stock return and house market return have moved in the same direction. From the second half of 2010, the upward trend has been reverted and the two asset returns exhibit a tendency to return back to its mean value by the end of the sample.

Figure 2.3: The dynamic correlations between stock and house returns



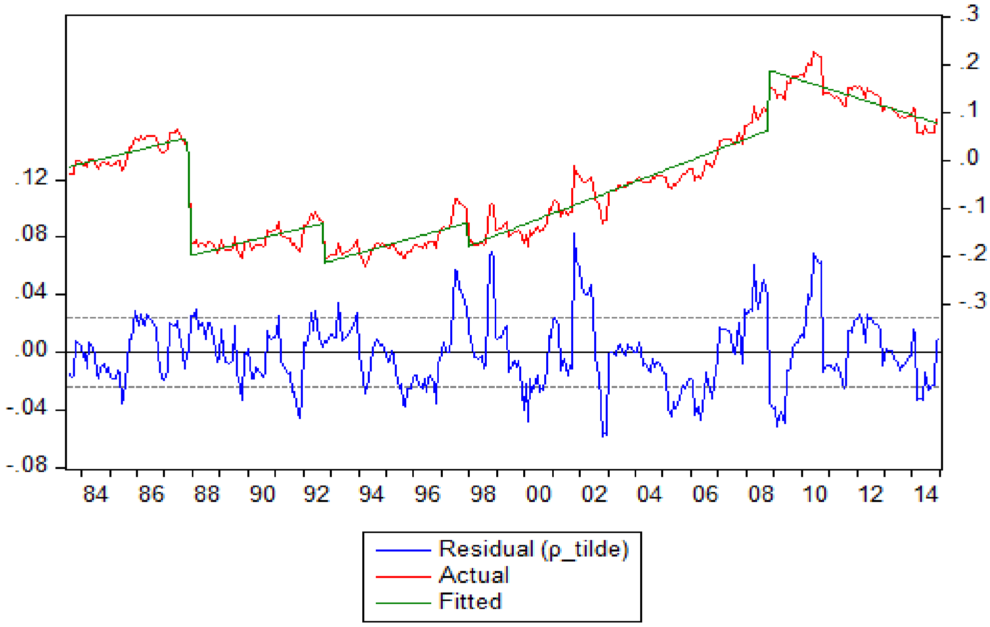

In a nutshell, the dynamic pattern of conditional correlations between the stock market return and the house market return displays significant fluctuations during the sample period with a mean reversion profile and high persistence. A natural question that arises is whether monetary policy developments could explain the variability of dynamic correlations between stock and house markets. In order to answer this question, we estimate a quantile model where the response variable is associated with predetermined monetary policy measures. If changes in dynamic correlations are positively affected by monetary regressors in upper tail percentiles, it means that an expansionary monetary policy will increase the variability between the stock and the house returns. On the other hand, if changes in dynamic correlations are negatively affected by monetary regressors in lower tail percentiles, it means that an expansionary monetary policy will decrease the variability between the stock and house returns. In order to measure the variability of stock and house returns, we regress initially the stationary correlation $\rho_t$ on a constant, a trend and two dummies indicated by the modified DF test, and estimate the regression with least squares using heteroskedasticity-autocorrelation consistent (HAC) standard errors. This filtering technique is similar to the approach proposed by Hamilton (2017) for a stationary variable. The Lagrange multiplier (LM) statistic of the KPSS unit root test with a constant and a trend is equal to 0.2588 which is greater than the asymptotic critical value of 0.2160 at 1% level, and thus the estimated residuals from this regression are not stationary. An inspection of the plotted residuals has indicated the presence of changing trends and the existence of an outlier at 1987M11. As a result, we have re-estimated the above regression with least squares with breaks using the procedure developed by Bai and Perron (2003).[5] This method permits up to five breaks and uses a trimming threshold of 15% for the data. The covariance matrix is estimated with HAC standard errors. The results which are presented in Table 2.4 show that the intercept and the trend are changing across breakpoints and are both highly significant. The breaks are located at the following dates: 1992M10, 1997M12 and 2008M11. The first break reflects the recession of 1992, where unemployment increased over 10%, and the events which followed the decision of the government to leave the exchange rate mechanism. The second break reflects the initial effects of the global financial crisis on the economy which started with the collapse of the Northen Rock in 2007, while the third break the recession of the economy as a result of the global credit crunch.

[5] The regression includes a third dummy $D3$ which takes 1 on the date 1987M11 and 0 elsewhere, and captures the presence of an outlier.

| **Table 2.4: Bai-Perron estimates with breaks of dynamic correlations** | | |
|---|---|---|
| Samples | Constant | Trend |
| 1983M07 - 1992M09 -- 111 obs | -0.023 ** (0.007) | 0.001 *** (0.0002) |
| 1992M10 - 1997M11 -- 62 obs | -0.127 ** (0.05) | 0.001 *** (0.0003) |
| 1997M12 - 2008M10 -- 131 obs | -0.268 *** (0.06) | 0.002 *** (0.0001) |
| 2008M11 - 2014M12 -- 74 obs | 0.902 *** (0.117) | -0.002 *** (0.0003) |
| *Non-Breaking Variables* | | |
| $D1$ | 0.237 *** (0.007) | |
| $D2$ | -0.247 *** (0.013) | |
| $D3$ | 0.199 *** (0.007) | |
| Adjusted $R$-squared = 0.96, SEE=0.02, Durbin-Watson stat =0.39 | | |
| *Sequential F-statistic determined breaks:* 3 | | |
| | $F$-statistic | *Scaled F*-statistic [CV] |
| 0 vs. 1 ** | 24.181 | 48.362 [11.47] |
| 1 vs. 2 ** | 43.937 | 87.873 [12.95] |
| 2 vs. 3 ** | 7.551 | 15.101 [14.03] |
| 3 vs. 4 | 2.250 | 4.499 [14.85] |
| *Break dates* | | |
| | Sequential | Repartition |
| 1 | 2007M12 | 1992M10 |
| 2 | 1992M10 | 1997M12 |
| 3 | 1997M12 | 2008M11 |
| **Notes**: Sample size: 384. ***, ** indicate significance at 1% and 5% respectively. Bai-Perron (2003) critical values (CV), obs=observations. | | |

Figure 2.4 plots the actual and fitted values of $\rho_t$, and the estimated residuals which measure the variability of dynamic correlations, denoted as $\widetilde{\rho}_t$. The LM statistic of KPSS unit root test is equal to 0.033 which is smaller than the asymptotic critical value of 0.4630 at 5% level, and thus $\widetilde{\rho}_t$ is a stationary process.

Figure 2.4: The variability of dynamic correlations




### 2.5.4. Quantile analysis

Before presenting and discussing the results from bivariate and multivariate analysis which span several pages, the conclusion which emerges from these results and answers the research question is that monetary policy measures in UK, such as M0, M4, and the policy interest rate, have substantial predictive content for the variability of dynamic correlations between stock and house markets returns. This evidence from the UK adds to the existing literature that stock and house markets constitute mechanisms though which monetary policy decisions affect the macroeconomic environment, and thus these facts should be taken into consideration when policy decisions are designed. In other words, the answer to the research question provides policy makers in the UK with valuable information about monetary policy developments in characterizing the stylized facts of dynamic co-movements of stock-house returns, and benefit investors in diversifying their portfolio strategies.

*Bivariate regression results*

Having extracted the variability of dynamic correlations, we proceed then to estimate the quantile regression (2.4.8). In this sub-section, we present and discuss the results from the bivariate model which regresses the response variable $\tilde{\rho}_t$ on a lagged value of its own and a monetary regressor measured as three-month and six-month moving averages. Tables 2.5-2.8 present quantile estimates at five intervals: $\tau \in \{0.05, 0.25, 0.50, 0.75, 0.95\}$. In all estimated regressions, the *Q*-statistics showed that the estimated residuals do not exhibit serial correlation at thirty lags for all specified intervals except for 95th percentile. The following remarks emerge from the bivariate analysis:

1. The autoregressive coefficient of the response variable varies across quantiles within a range from 0.71 to 0.92 indicating that the dynamics of the variability of dynamic correlations exhibit an asymmetry in the persistence profile. Particularly, in most cases the value of the autoregressive coefficient increases as we move from lower quantiles to median where it achieves its highest value, and then it falls as we move from median to higher quantiles. Thus, positive and negative shocks have less persistent effects on the variability of dynamic correlations in the lower and upper quantiles. The null hypothesis that the autoregressive coefficient is equal to one against the alternative hypothesis that the coefficient is less than one can be tested using the covariate augmented DF (CADF) *t*-test. This test was applied to equation (2.4.8) with the three-month moving average of the broad money M4. The autoregressive coefficients are 0.88 ($\tau$=0.95), 0.915 ($\tau$=0.5) and 8.87 ($\tau$=0.25). The CADF *t*-statistic was equal to -4.369 ($\tau$=0.95), -4.126 ($\tau$=0.5) and -4.45 ($\tau$=0.25). The critical values for the demeaned model at 5% significance level are -3.39 and -2.72 for $\delta^2$=0.9 and 0.70 respectively. Thus, the test rejects the null hypothesis of a unit root at the three specified quantiles. Given that the estimated autoregressive

coefficients in our bivariate and multivariate models have similar size as above, we have not proceeded further with unit root tests in other models.

**2.** The three-month and six-month moving averages of money measures have a statistically significant impact on current variability of dynamic correlations, and the sign of their impact is positive in the higher quantiles (75th and 95th percentiles) and negative in the lower quantiles (5th and 25th percentiles). This asymmetric effect implies that when changes in dynamic correlations are positive, an increase in the money supply and its inside component will further increase the variability of correlations between stock and house returns, thus increasing the possibility of a boom episode. On the other hand, when changes in dynamic correlations are negative, an increase in the money supply and its inside component will further decrease the variability of correlations between stock and house returns, thus increasing the possibility of a bust episode. This finding shows that predetermined movements in the broad and narrow money, and inside money are important to understand future movements in the variability of dynamic correlations between the stock market return and the house market return, given the information provided by its own past history.

**3.** The three-month and six-month moving averages of changes in the official rate have a statistically significant impact on current variability of correlations, and the sign of its impact is negative in the higher quantiles (75th and 95th percentiles) and positive in the lower quantiles (5th and 25th percentiles). This asymmetric interest rate effect implies that when variability is positive, a fall in the official rate will further increase the variability of stock-house returns, thus increasing the possibility of a boom episode. On the other hand, when variability is negative, a fall in the interest rate will further decrease the variability of stock-house returns, thus increasing the possibility of a bust episode. This finding shows that predetermined movements in the official interest rate have information content that is useful to understand future movements in the variability of dynamic correlations between the stock market return and the house market return, given the information provided by the variability's past history.

| **Table 2.5: Quantile estimates from bivariate model: Broad money M4** | | | | | | |
|---|---|---|---|---|---|---|
| | | *Quantiles* | | | | |
| | Regressors | 0.05 | 0.25 | 0.5 | 0.75 | 0.95 |
| **MA(3)** | $\tilde{\rho}_{t-1}$ | 0.793***<br>(0.033) | 0.868***<br>(0.023) | 0.915***<br>(0.020) | 0.915***<br>(0.022) | 0.899***<br>(0.023) |
| | $\Delta\overline{m}$ | -0.017***<br>(0.002) | -0.005***<br>(0.001) | -0.001<br>(0.005) | 0.004***<br>(0.001) | 0.016***<br>(0.001) |
| **MA(6)** | $\tilde{\rho}_{t-1}$ | 0.752***<br>(0.05) | 0.877***<br>(0.033) | 0.916***<br>(0.02) | 0.893***<br>(0.028) | 0.847***<br>(0.032) |
| | $\Delta\overline{m}$ | -0.02***<br>(0.003) | -0.006***<br>(0.001) | -0.001<br>(0.001) | 0.005***<br>(0.001) | 0.02***<br>(0.002) |

**Notes:** Sample size: 384. The estimated model has the form: $Q_{\tilde{\rho}_t}\left(\tau \middle| \tilde{\rho}_{t-1}, \Delta\overline{m}\right) = \left(1, \tilde{\rho}_{t-1}, \Delta\overline{m}\right)' \vartheta(\tau)$. MA(*k*)= *k*-month moving average of monthly growth rates of broad money M4 $(\Delta\overline{m})$. Numbers in parentheses are SEs. *** indicates significance at 1% level.

| **Table 2.6: Quantile estimates from bivariate model: Narrow money M0** | | | | | | |
|---|---|---|---|---|---|---|
| | | *Quantiles* | | | | |
| | Regressors | 0.05 | 0.25 | 0.5 | 0.75 | 0.95 |
| **MA(3)** | $\tilde{\rho}_{t-1}$ | 0.768***<br>(0.053) | 0.882***<br>(0.033) | 0.915***<br>(0.021) | 0.881***<br>(0.026) | -0.038***<br>(0.003) |
| | $\Delta\overline{m}$ | -0.038***<br>(0.003) | -0.011***<br>(0.002) | -0.001<br>(0.001) | 0.007***<br>(0.001) | 0.054***<br>(0.001) |
| **MA(6)** | $\tilde{\rho}_{t-1}$ | 0.736***<br>(0.05) | 0.882***<br>(0.031) | 0.915***<br>(0.021) | 0.862***<br>(0.028) | 0.843***<br>(0.068) |
| | $\Delta\overline{m}$ | -0.041***<br>(0.004) | -0.012***<br>(0.002) | -0.001<br>(0.001) | 0.009***<br>(0.001) | 0.050***<br>(0.008) |

**Notes:** Sample size: 384. MA(*k*)= *k*-month moving average of monthly growth rates of narrow money M0 $(\Delta\overline{m})$. *** indicates significance at 1% level.

| Table 2.7: Quantile estimates from bivariate model: Inside money (IM=M4-M0) | | | | | | |
|---|---|---|---|---|---|---|
| | | *Quantiles* | | | | |
| | Regressors | 0.05 | 0.25 | 0.5 | 0.75 | 0.95 |
| **MA(3)** | $\tilde{\rho}_{t-1}$ | 0.720*** (0.040) | 0.792*** (0.033) | 0.915*** (0.021) | 0.920*** (0.022) | 0.895*** (0.022) |
| | $\Delta\overline{m}$ | -0.020*** (0.002) | -0.017*** (0.002) | -0.001 (0.001) | 0.004*** (0.001) | 0.015*** (0.001) |
| **MA(6)** | $\tilde{\rho}_{t-1}$ | 0.756*** (0.050) | 0.874*** (0.033) | 0.916*** (0.020) | 0.893*** (0.027) | 0.851*** (0.031) |
| | $\Delta\overline{m}$ | -0.019*** (0.003) | -0.006*** (0.001) | -0.001 (0.001) | 0.004*** (0.001) | 0.019*** (0.001) |

**Notes:** Sample size: 384. MA(*k*)= *k*-month moving average of monthly growth rates of inside money IM $(\Delta\overline{m})$. *** indicates significance at 1% level.

| Table 2.8: Quantile estimates from bivariate model: Official interest rate | | | | | | |
|---|---|---|---|---|---|---|
| | | *Quantiles* | | | | |
| | Regressors | 0.05 | 0.25 | 0.5 | 0.75 | 0.95 |
| **MA(3)** | $\tilde{\rho}_{t-1}$ | 0.923*** (0.009) | 0.917*** (0.020) | 0.911*** (0.021) | 0.906*** (0.018) | 0.908*** (0.009) |
| | $\Delta\overline{R}$ | 0.002*** (0.001) | 0.001 (0.001) | -5.8E-05 (0.002) | -0.001 (0.001) | -0.002*** (0.001) |
| **MA(6)** | $\tilde{\rho}_{t-1}$ | 0.922*** (0.009) | 0.918*** (0.017) | 0.911*** (0.021) | 0.905*** (0.018) | 0.914*** (0.010) |
| | $\Delta\overline{R}$ | 0.003*** (0.001) | 0.002 (0.002) | -4.8E-05 (0.002) | -0.002 (0.002) | -0.008*** (0.001) |

**Notes:** Sample size: 384. The estimated model has the form: $Q_{\tilde{\rho}_t}\left(\tau \middle| \tilde{\rho}_{t-1}, \Delta\overline{R}\right) = \left(1, \tilde{\rho}_{t-1}, \Delta\overline{R}\right)' \vartheta(\tau)$. MA(*k*)= *k*-month moving average of monthly changes in the official interest rate $(\Delta\overline{R})$. *** indicates significance at 1% level.

Figures 2.5-2.8 plot the quantile estimates with the 95% confidence bands from the bivariate model. These plots provide a clear-cut picture of the importance of monetary measures as indicators of changes in dynamic correlations. The fact that the quantile estimates cross the median estimates suggest that an OLS analysis of the issue at hand is misleading.

Figure 2.5: Quantile estimates from the bivariate model (Broad Money M4)

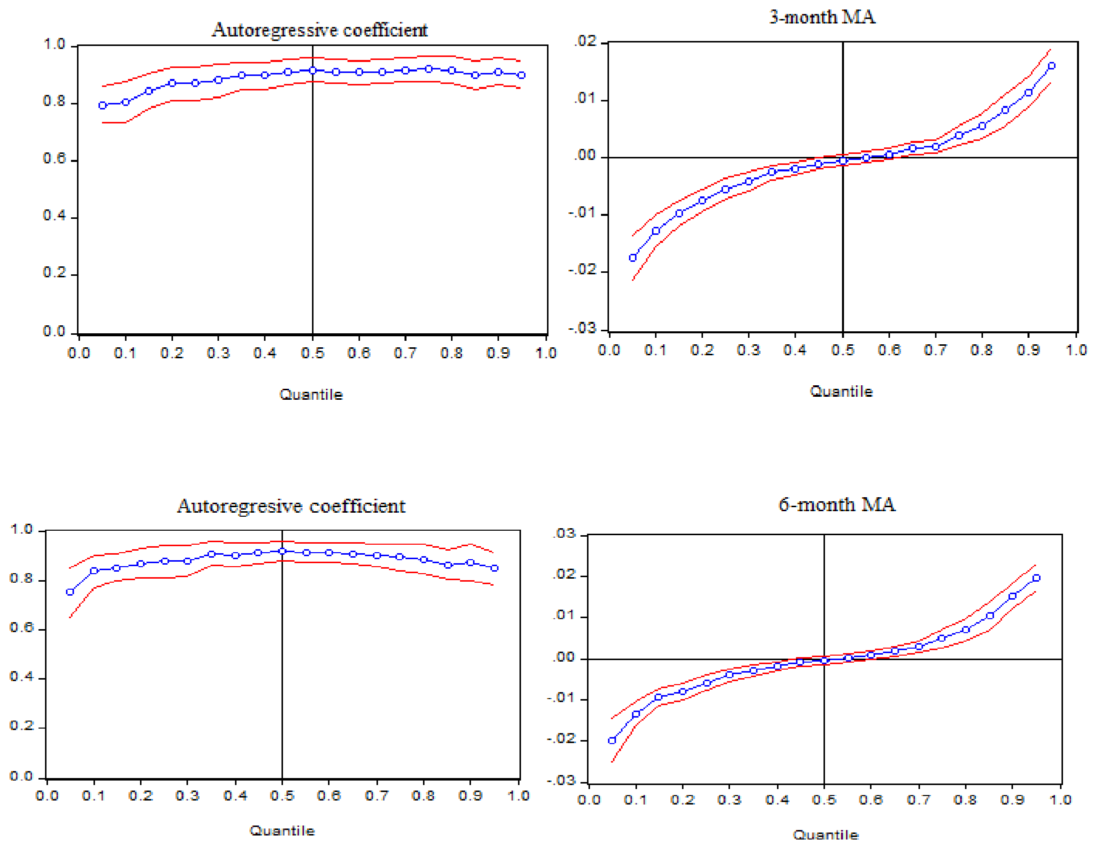


Figure 2.6: Quantile estimates from the bivariate model (Narrow Money M0)

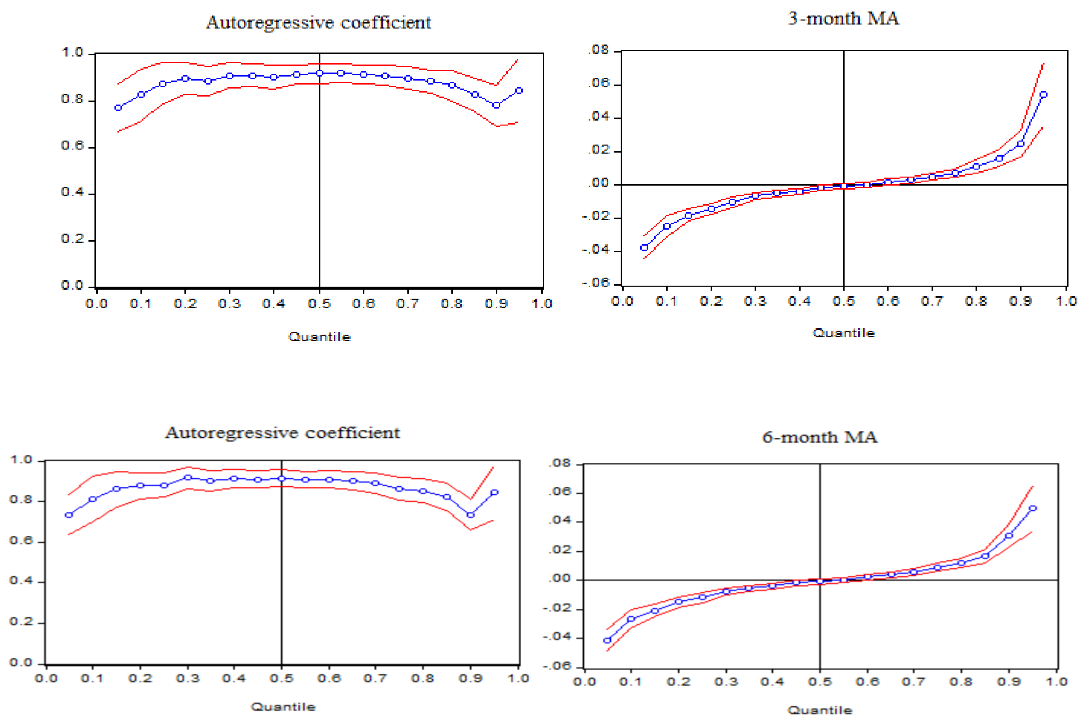



Figure 2.7: Quantile estimates from the bivariate model (Inside Money)

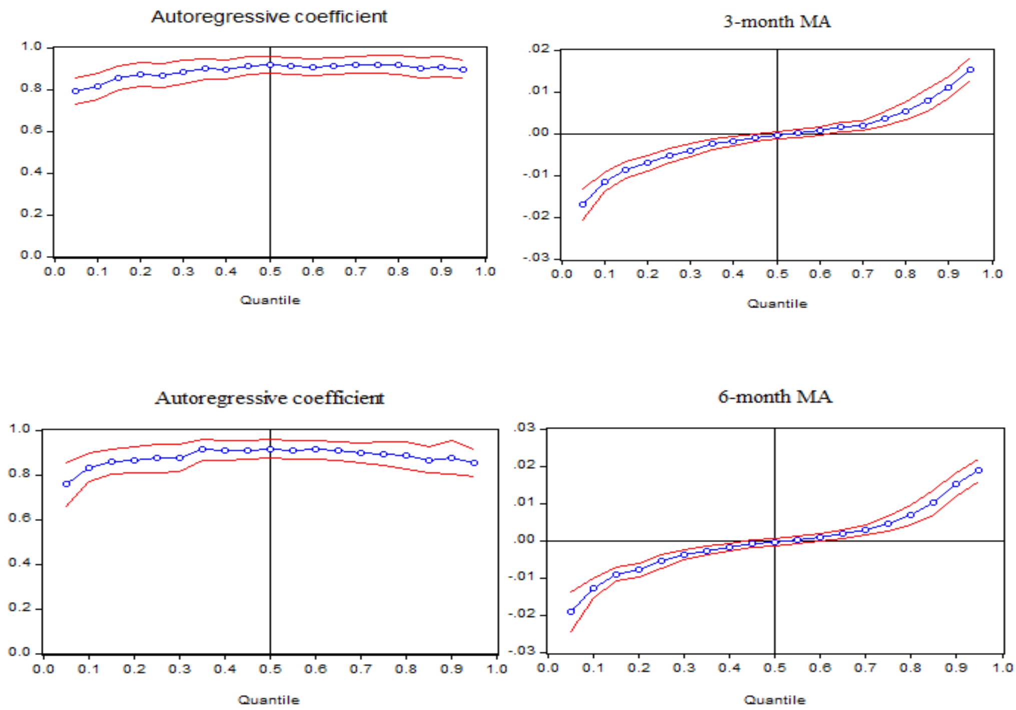
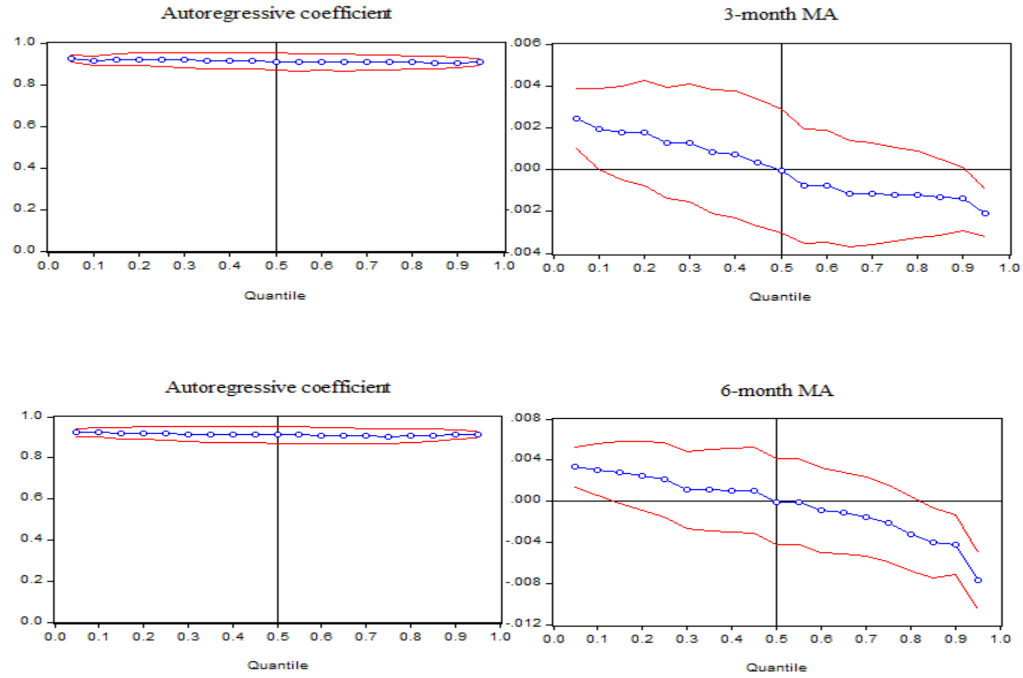


Figure 2.8: Quantile estimates from the bivariate model (Official interest rate)




*Multivariate regression results with control variables*

In order to check the robustness of the aforementioned results, it is important to control for the impact that other predictors considered in the literature may have on the variability of dynamic correlations of stock-house returns. Thus, we have estimated a multivariate version of equation (2.4.8) which, except from monetary variables measured as three-month and six-month moving averages, also includes six macroeconomic variables measured as three-month and six-month moving averages. The control variables which we utilize as predictors are the following: the price level, the industrial production, the unemployment rate, the 10-year government bond interest rate, the total lending, and the effective exchange rate.[6] The stationarity of the control variables is attained by taking the first-difference of the logs of prices, industrial production, total lending, and exchange rate, and the first-difference of unemployment rate and long-term interest rate. We have constructed three-month and six month moving averages of monthly growth rates of the former four variables, and three-month and six-month moving averages of monthly changes of the latter two variables.

Figures 2.9-2.12 plot the quantile estimates with the 95% confidence bands from the multivariate model using three-month moving averages of monetary measures and control variables. On inspection we observe that all monetary measures retain their predictive content at lower and higher percentiles, but only M0 has a persistent impact across various quantiles. The signs of quantile estimates of monetary measures from the multivariate model are the same as the signs from the bivariate model. There are noticeable differences across quantiles with respect to the included control variables. The inflation rate impacts the variability positively in the higher quantile and negatively in the lower quantile, thus increasing the possibility of boom and bust episodes, respectively. The change in the unemployment rate seems to be relevant for $\tau>0.60$ with a negative sign, indicating that a fall in unemployment rate predicts an increase in the variability of dynamic correlations. The change in the long-term rate, in turn, has a positive effect on the variability in the lower tail and a negative impact in the upper tail, while the growth rate of total lending is negatively associated with the variability in the lower tail.

[6] The control variables have been obtained from the FRED St. Louis. The total lending refers to monetary financial institutions' sterling net lending to private sector. The seasonal adjustment of the consumer price level and the 10-year bond yield applied the Census X12 method.

Figure 2.9: Quantile estimates from multivariate model (3-month MA of M4)

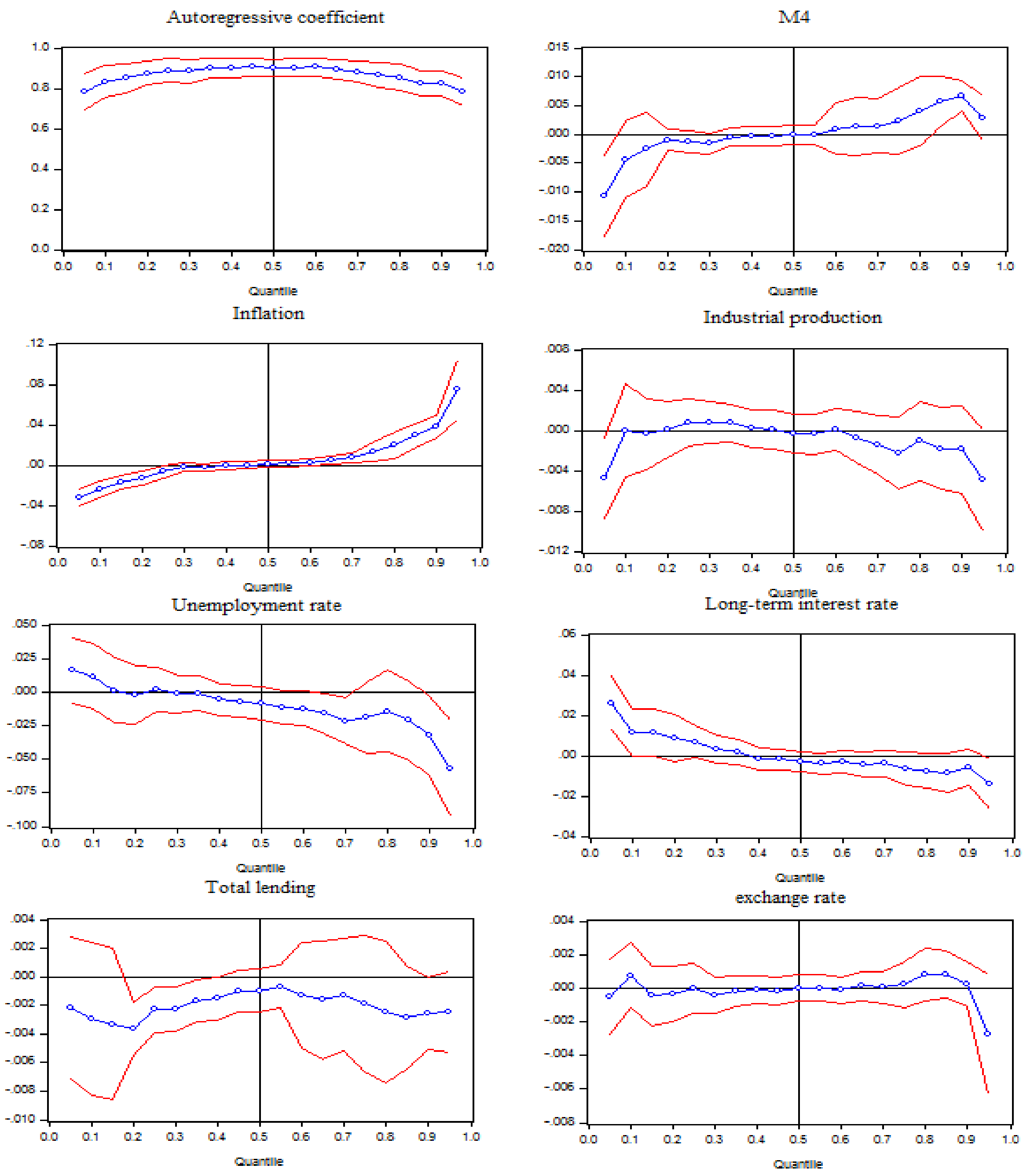


Figure 2.10: Quantile estimates from multivariate model(3-month MA of M0)

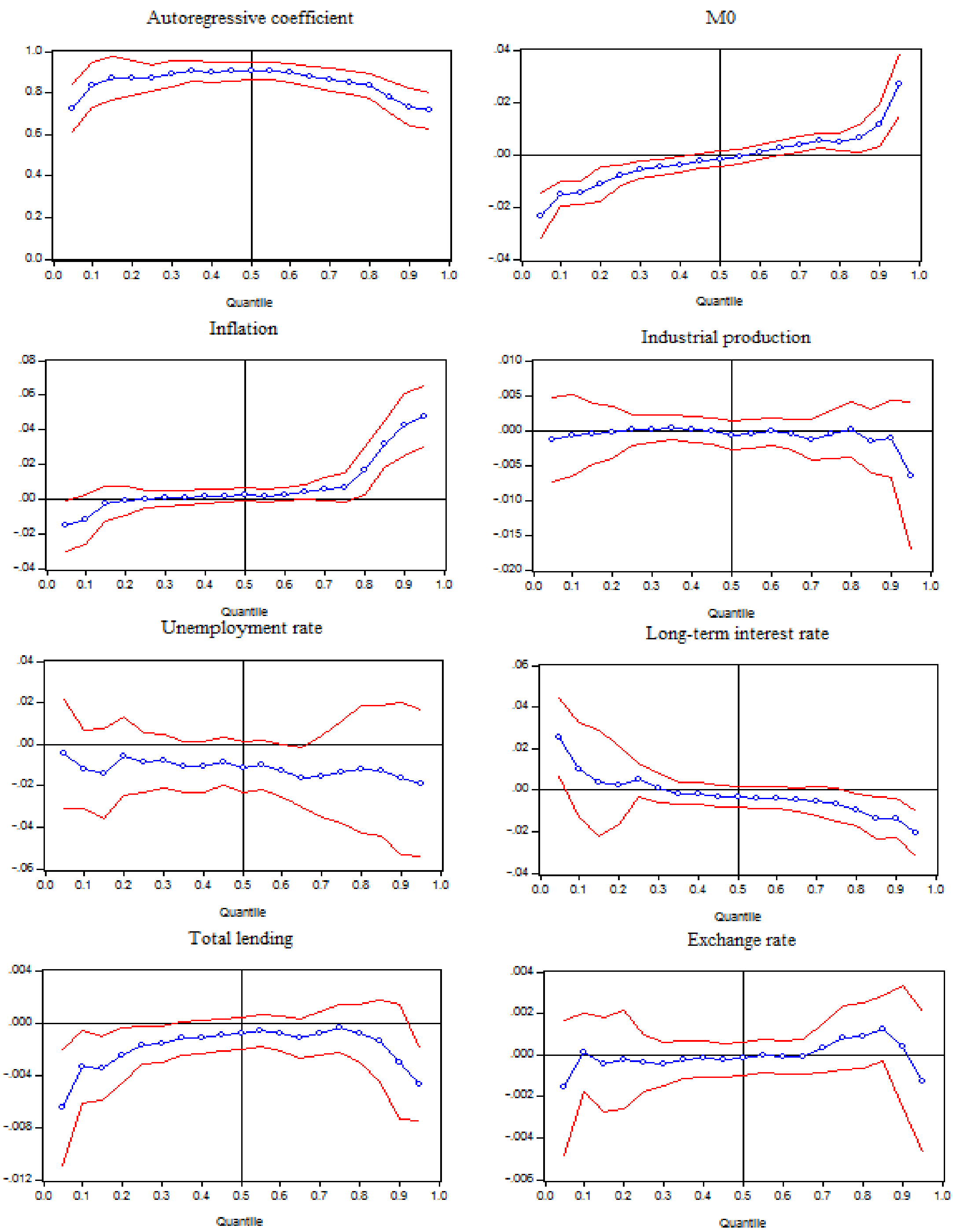


Figure 2.11: Quantile estimates from multivariate model(3-month MA of IM)

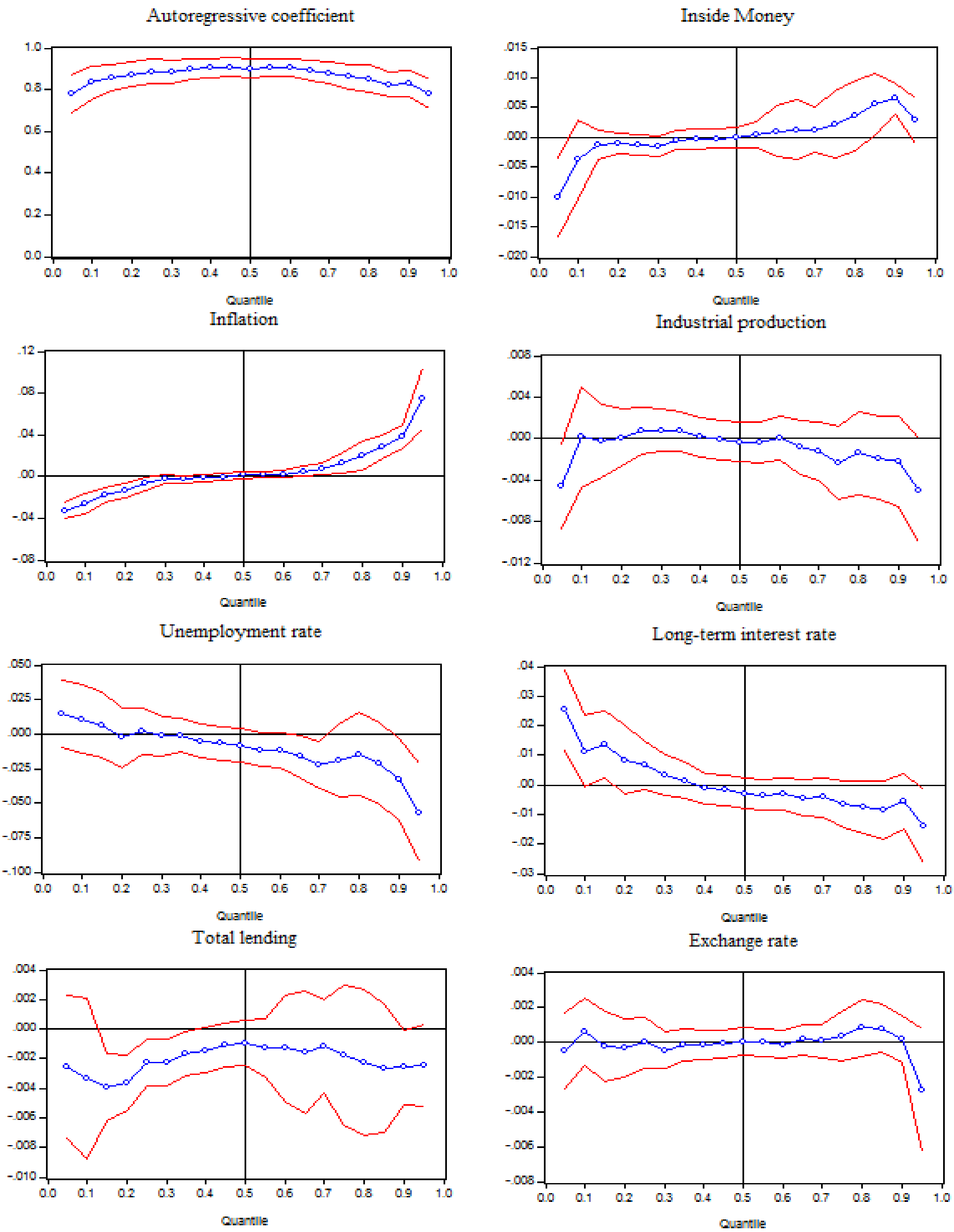
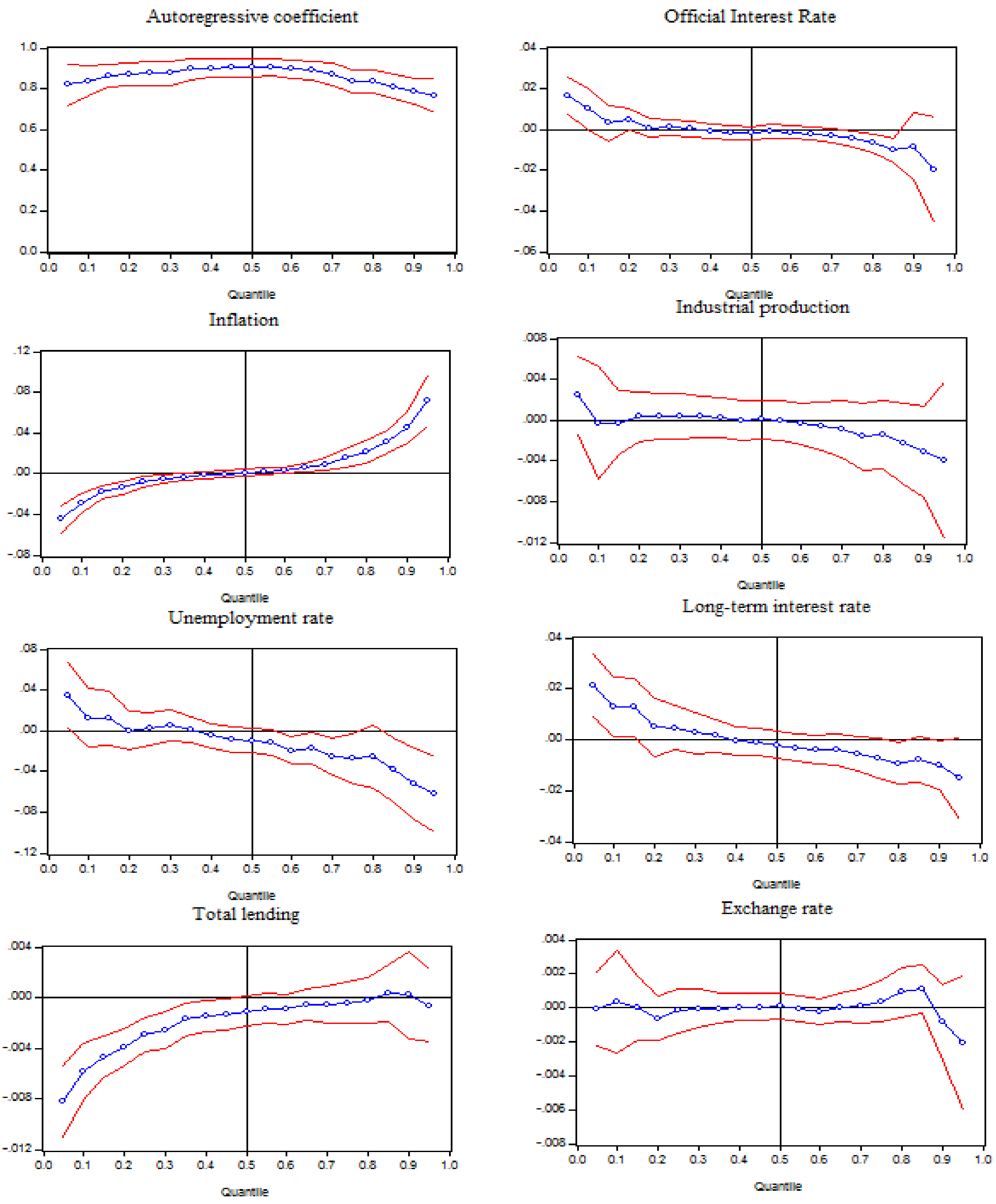


Figure 2.12: Quantile estimates from multivariate model (3-month MA of official interest rate)



Figures 2.13-2.16 plot the quantile estimates and the 95% confidence bands from the multivariate model using six-month moving averages of monetary measures and control variables. The figures suggest that monetary measures have predictive content for understanding changes in dynamic correlations, given the information content of other predictors and variability's past history.

Figure 2.13: Quantile estimates from multivariate model(6-month MA of M4)

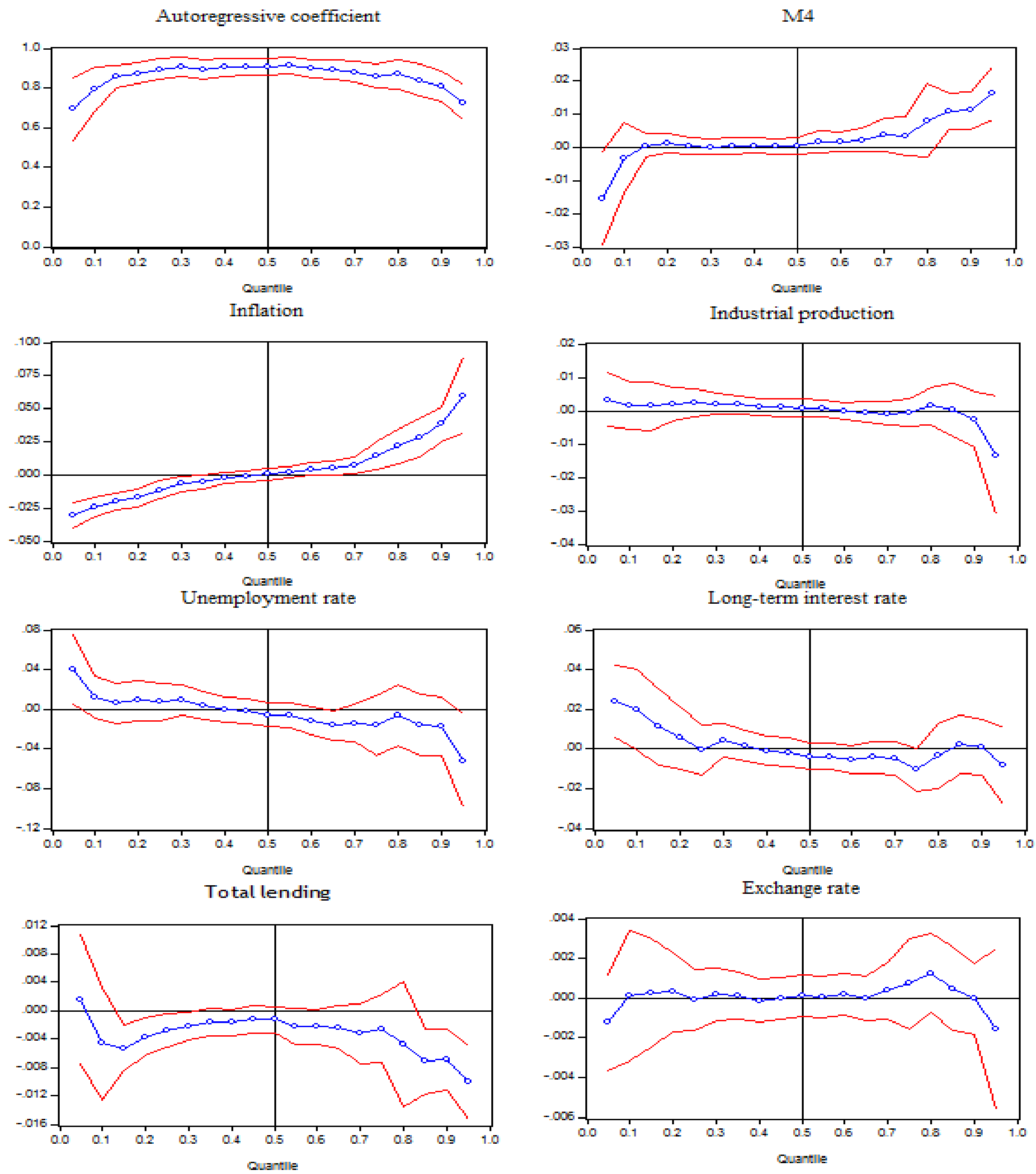


Figure 2.14: Quantile estimates from multivariate model(6-month MA of M0)

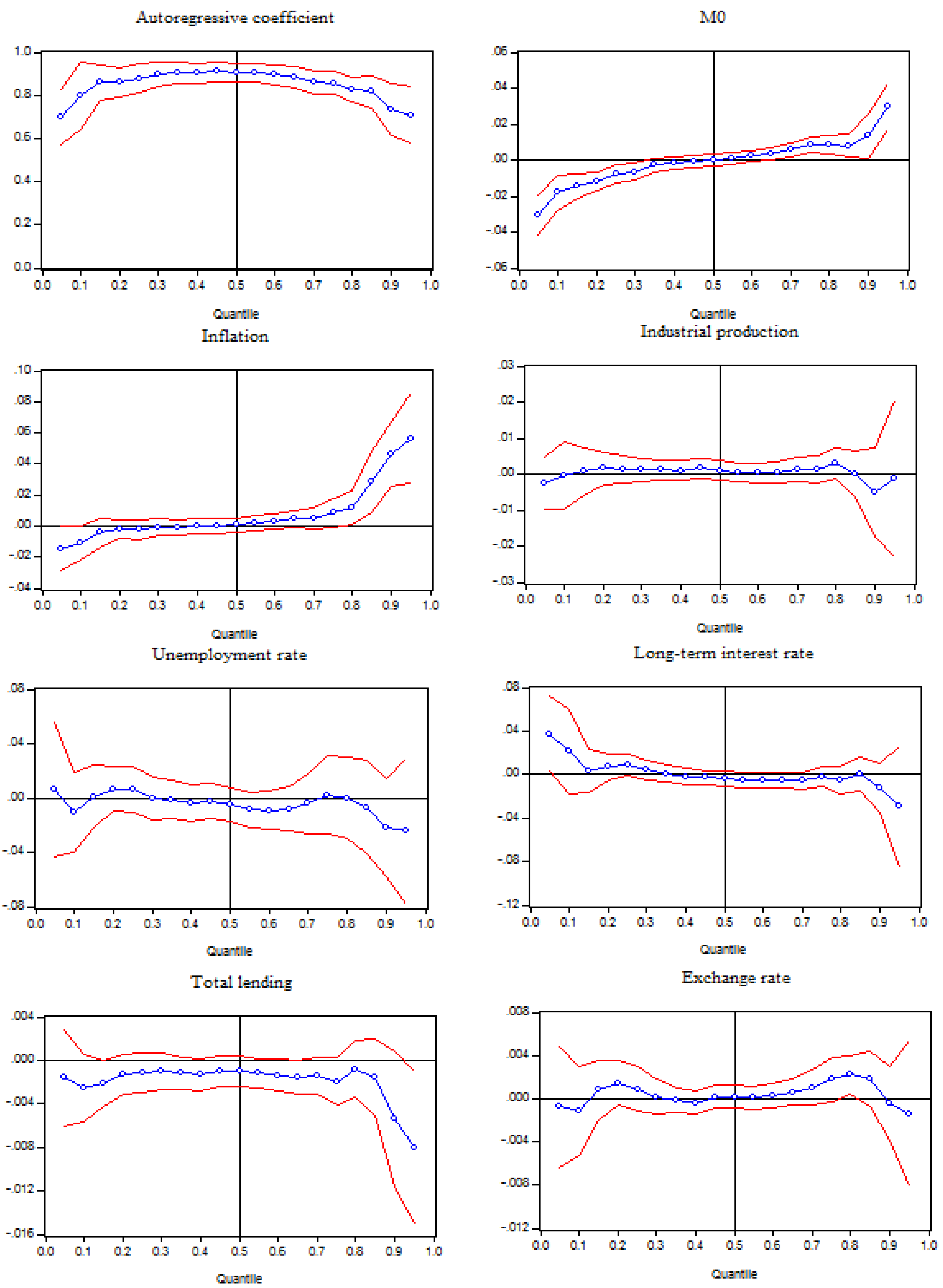


Figure 2.15: Quantile estimates from multivariate model(6-month MA of IM)

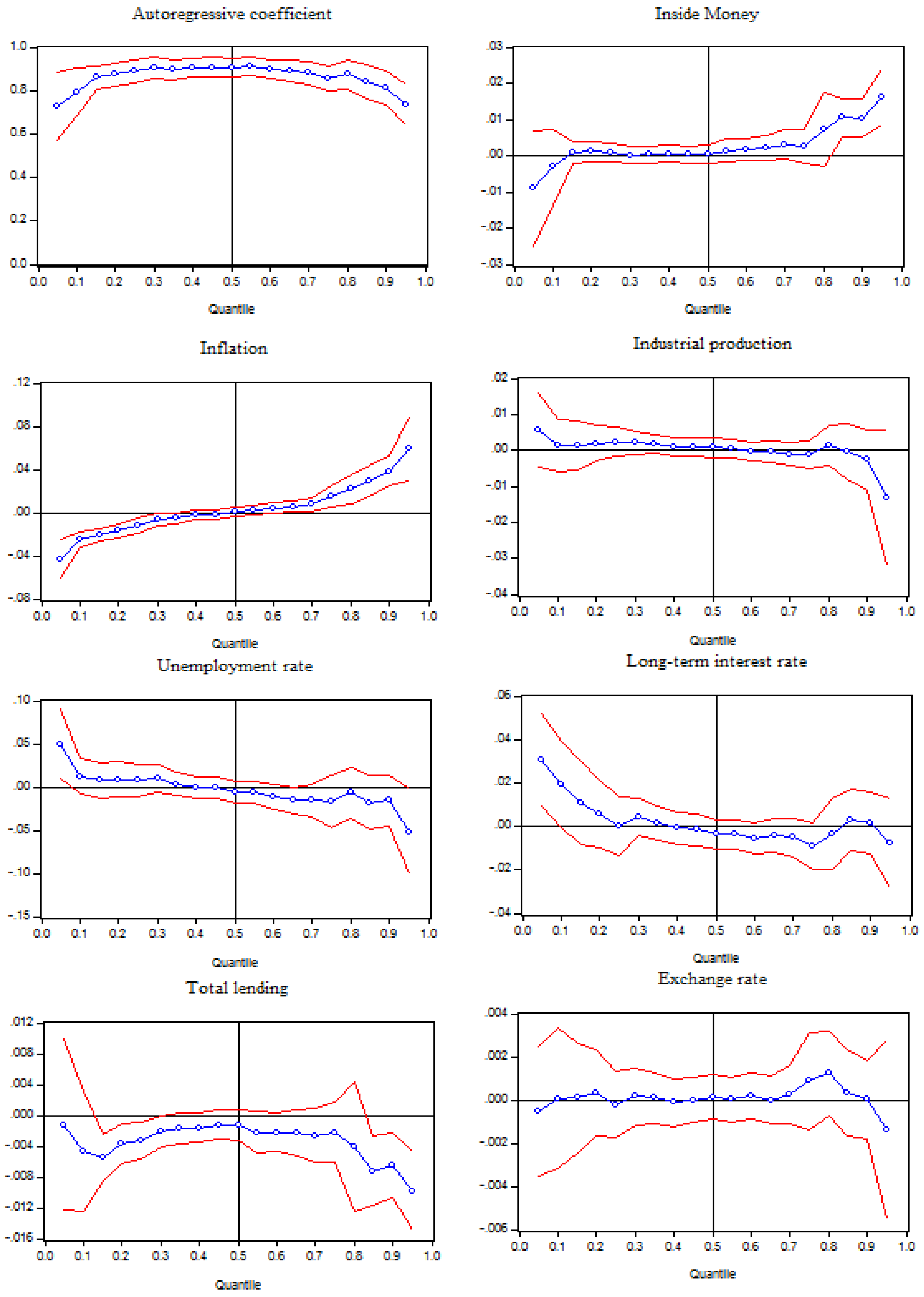
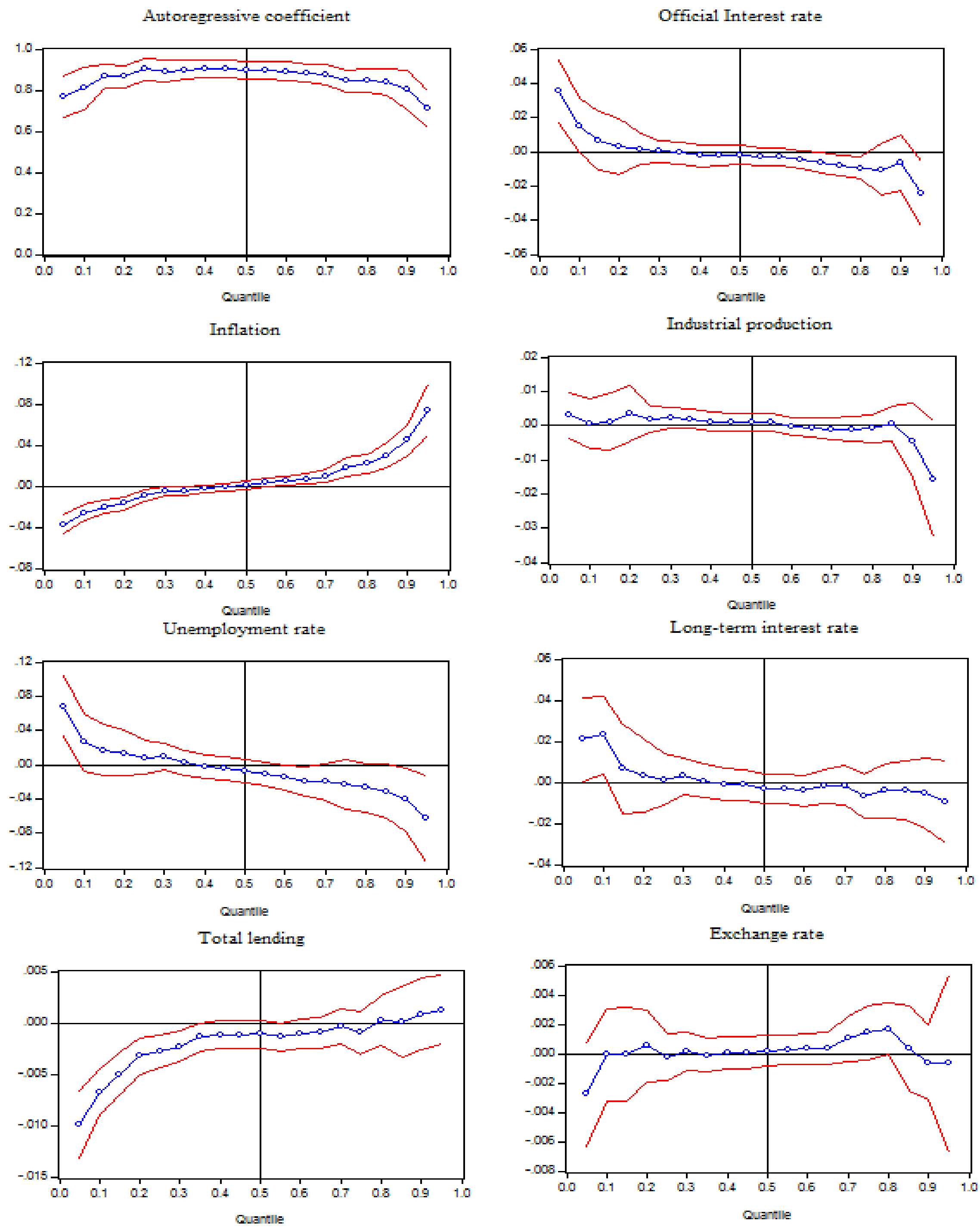


Figure 2.16: Quantile estimates from multivariate model(6-month MA of official interest rate)



### *Multivariate regression results with principal components*

The existence of possible correlation between the six macroeconomic variables diminishes seriously their significance to act as additional predictors in the quantile regression. To this end, we use a PCA to extract a lower dimensional space of linear combinations (the so-called principal components) between the six control variables which are uncorrelated from each other and explain the bulk of the variance of the original data. The first principal component is a linear combination of the six variables which accounts for the maximum variance of the data, the second principal component is a linear combination of the six variables which accounts for as much as possible of the remaining variance of the data and has the property that it is not correlated with the first component, and etc. In each principal component, the weights which are multiplied by each one of the six variables, constitute the eigenvectors (loadings) of the variance-covariance matrix of the original data. The proportion of variation explained by the first component is the eigenvalue for this component divided by the sum of the eigenvalues of all components (total variation). The cumulative proportion explained by the first and second component is the sum of their eigenvalues divided by the sum of the eigenvalues of all components. The selected principal components will be utilized as additional indicators in the quantile regression. Table 2.9 reports the PCA which is performed through the correlation matrix.

| Table 2.9: Principal component analysis using correlations | | | | | | |
|---|---|---|---|---|---|---|
| Eigenvalues: (Sum = 6, Average = 1) | | | | | | |
| | | | | Cumulative | Cumulative | |
| Number | Value | Difference | Proportion | Value | Proportion | |
| 1 | 1.291 | 0.123 | 0.217 | 1.291 | 0.217 | |
| 2 | 1.177 | 0.151 | 0.196 | 2.477 | 0.413 | |
| 3 | 1.026 | 0.042 | 0.171 | 3.503 | 0.584 | |
| 4 | 0.985 | 0.080 | 0.164 | 4.488 | 0.748 | |
| 5 | 0.904 | 0.296 | 0.151 | 5.392 | 0.899 | |
| 6 | 0.608 | --- | 0.101 | 6.000 | 1.000 | |
| Eigenvectors (loadings): | | | | | | |
| Variable | PC 1 | PC 2 | PC 3 | PC 4 | PC 5 | PC 6 |
| $\Delta p$ | 0.066 | 0.584 | 0.638 | 0.254 | -0.025 | -0.427 |
| $\Delta y$ | 0.499 | 0.039 | -0.035 | 0.198 | 0.829 | 0.147 |
| $\Delta u$ | -0.668 | 0.156 | 0.320 | 0.015 | 0.301 | 0.580 |
| $\Delta r$ | 0.432 | 0.264 | 0.250 | -0.708 | -0.162 | 0.391 |
| $\Delta tl$ | 0.132 | 0.587 | -0.467 | 0.420 | -0.296 | 0.394 |
| $\Delta e$ | 0.312 | -0.466 | 0.457 | 0.467 | -0.327 | 0.381 |

**Notes:** Sample size: 384. $\Delta p$, $\Delta y$, $\Delta u$, $\Delta r$, $\Delta tl$, $\Delta e$ denote, respectively, inflation, industrial production growth rate, change in unemployment rate, change in the long-term interest rate, growth rate of total lending, growth rate of exchange rate.

The upper panel shows the eigenvalues of the six possible components, the proportion explained by each component, and the cumulative proportion. Our purpose is to extract a small number of principal components which explains a large cumulative proportion of the total variation of the data. The first three components with eigenvalues greater than one explain 58% of the total variation.[7] Proceeding to the fourth component which explains 16% of the total variation, the cumulative proportion increases to about 75%. We regard this proportion excessively large, and thus we select the first four principal components, which are presented in the lower panel. In order to give an economic interpretation of the selected principal components, we look for control variables with loadings greater than 0.5 in which case these variables are strongly correlated with each component. The first principal

[7] The selection of principal components with eigenvalues greater than one is based on the Kaiser-Guttman criterion.

component is positively correlated with the growth rate of the industrial production and negatively correlated with the change in the unemployment rate. This component which increases with real output and decreases with unemployment rate reflects the Okun's Law. The second principal component is positively correlated with inflation and the growth rate of lending. Given that credit growth is lagging behind the recovery in economic activity, and that economic recovery and inflation are procyclical over the business cycle, the second component can be viewed as a measure of economic recovery. The third component increases with inflation and depicts the inflationary effect on the economy, and finally the fourth component which increases with a fall in the long-term interest rate can be viewed as reflecting the expectations theory of the term structure which relates the long-term yield to the expected one-period discount rates and consequently to divided-price ratio.

In turn, we construct the three-month and six-month moving averages of the four principal component and used them as additional predictors in a multivariate version of Equation (2.4.5), which also includes a lagged value of the variability of dynamic correlations and a monetary variable measured as three-month and six-month moving average. Figures 2.17-2.24 plot the quantile estimates and the 95% confidence bands of the multivariate model with principal components.

Figure 2.17: Quantile estimates from the multivariate model with principal components (3-month MA of M4)

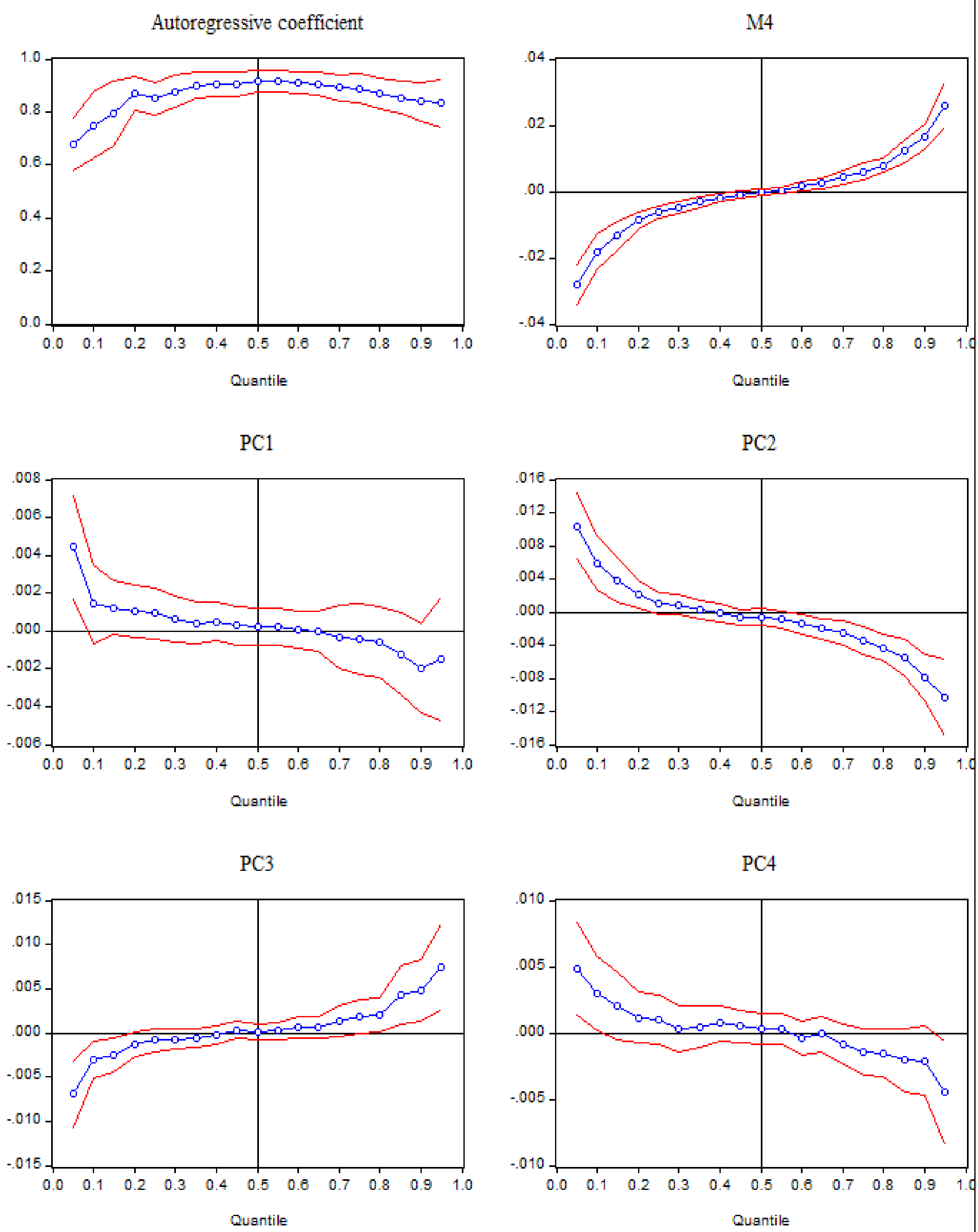


Figure 2.18: Quantile estimates from the multivariate model with principal components (3-month MA of M0)

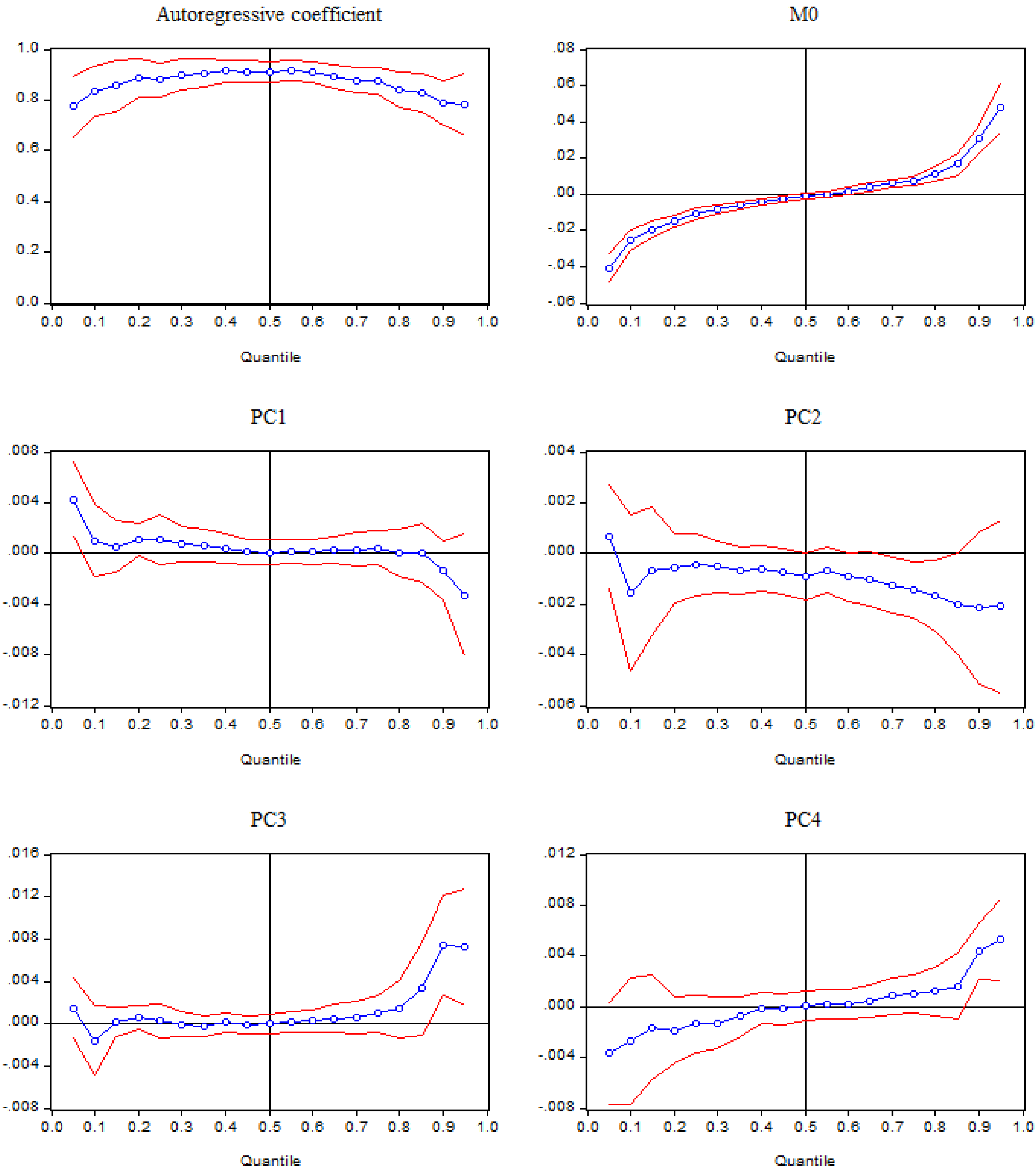


Figure 2.19: Quantile estimates from the multivariate model with principal components (3-month MA of IM)

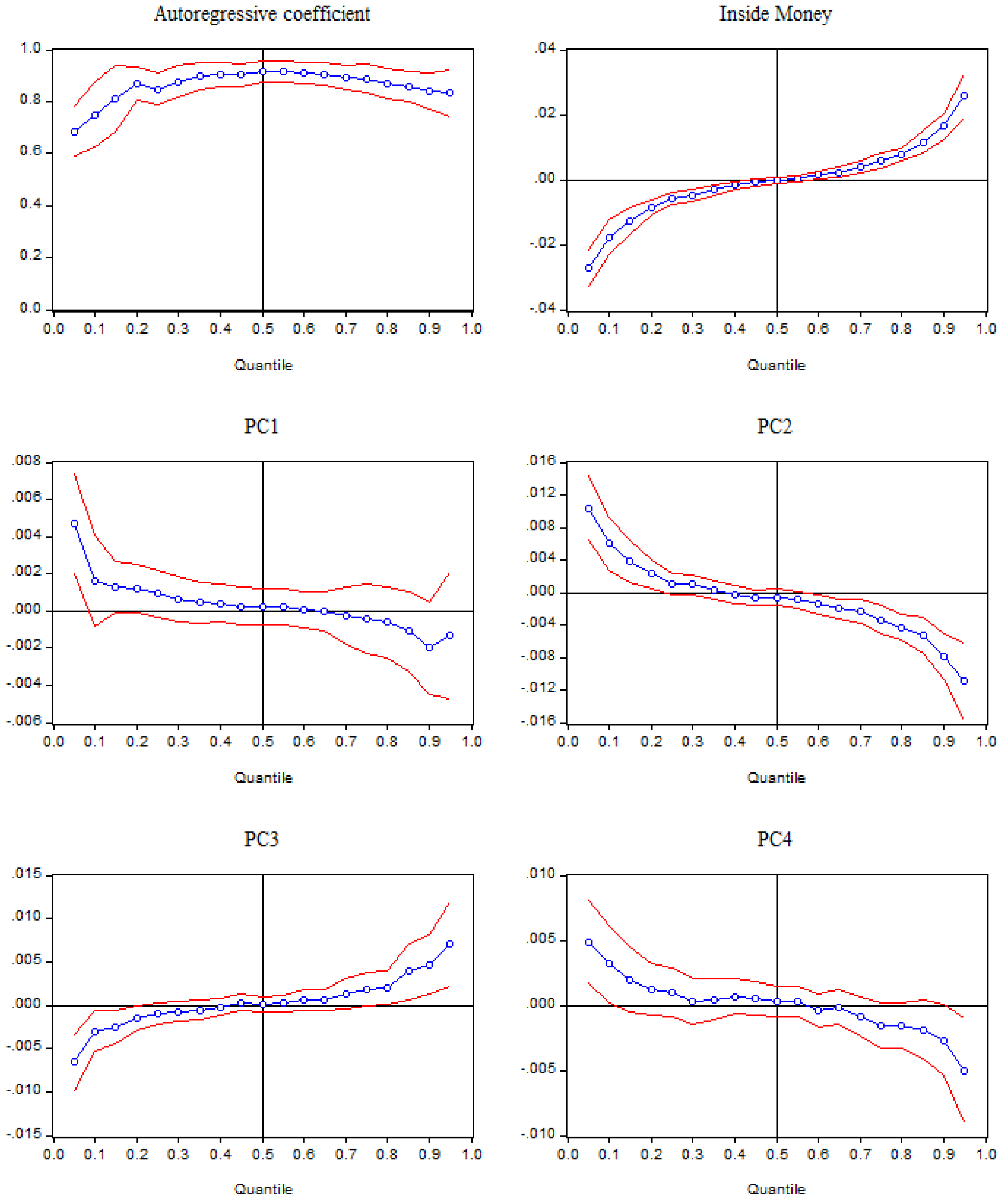


Figure 2.20: Quantile estimates from the multivariate model with principal components (3-month MA of official interest rate)

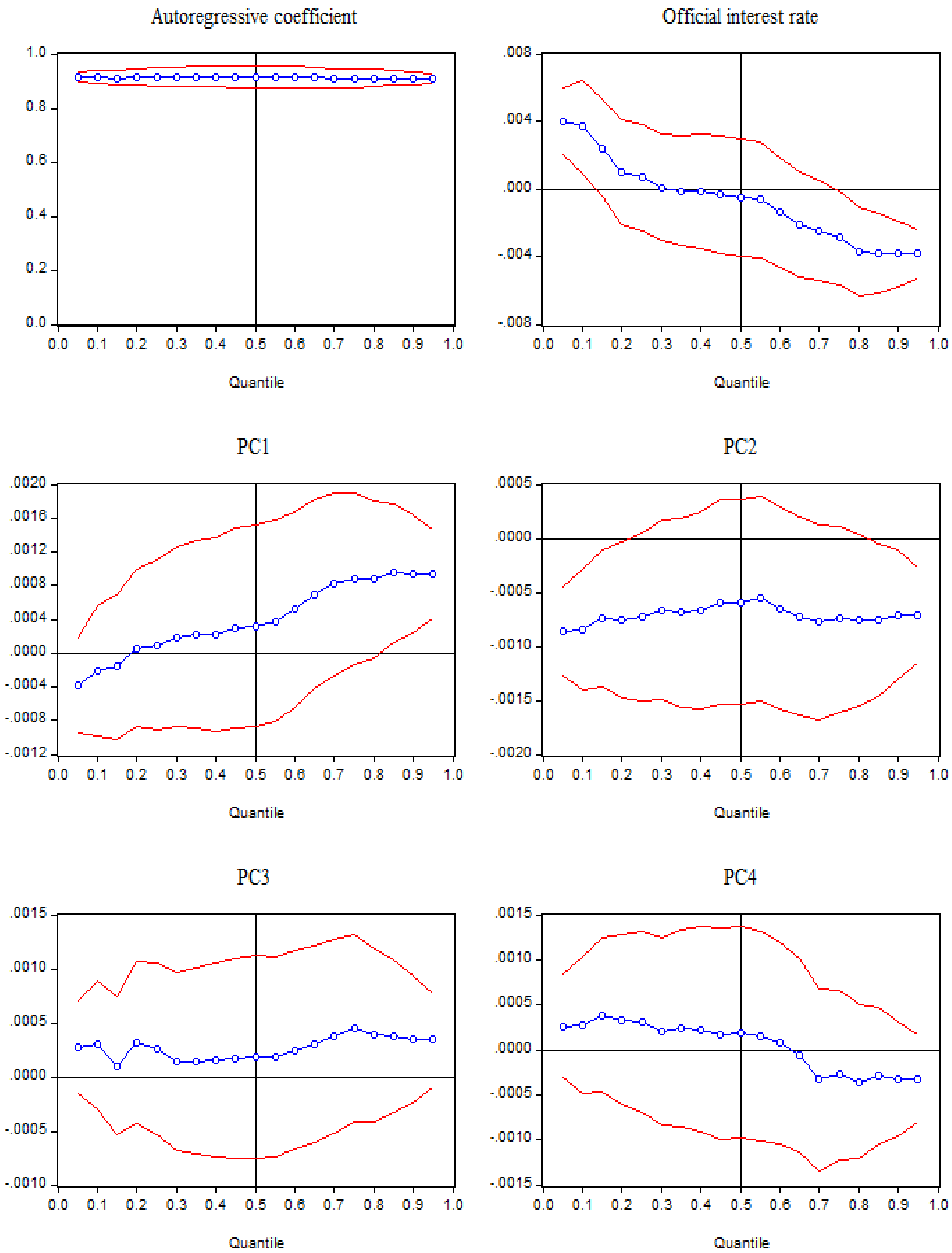


Figure 2.21: Quantile estimates from the multivariate model with principal components (6-month MA of M4)

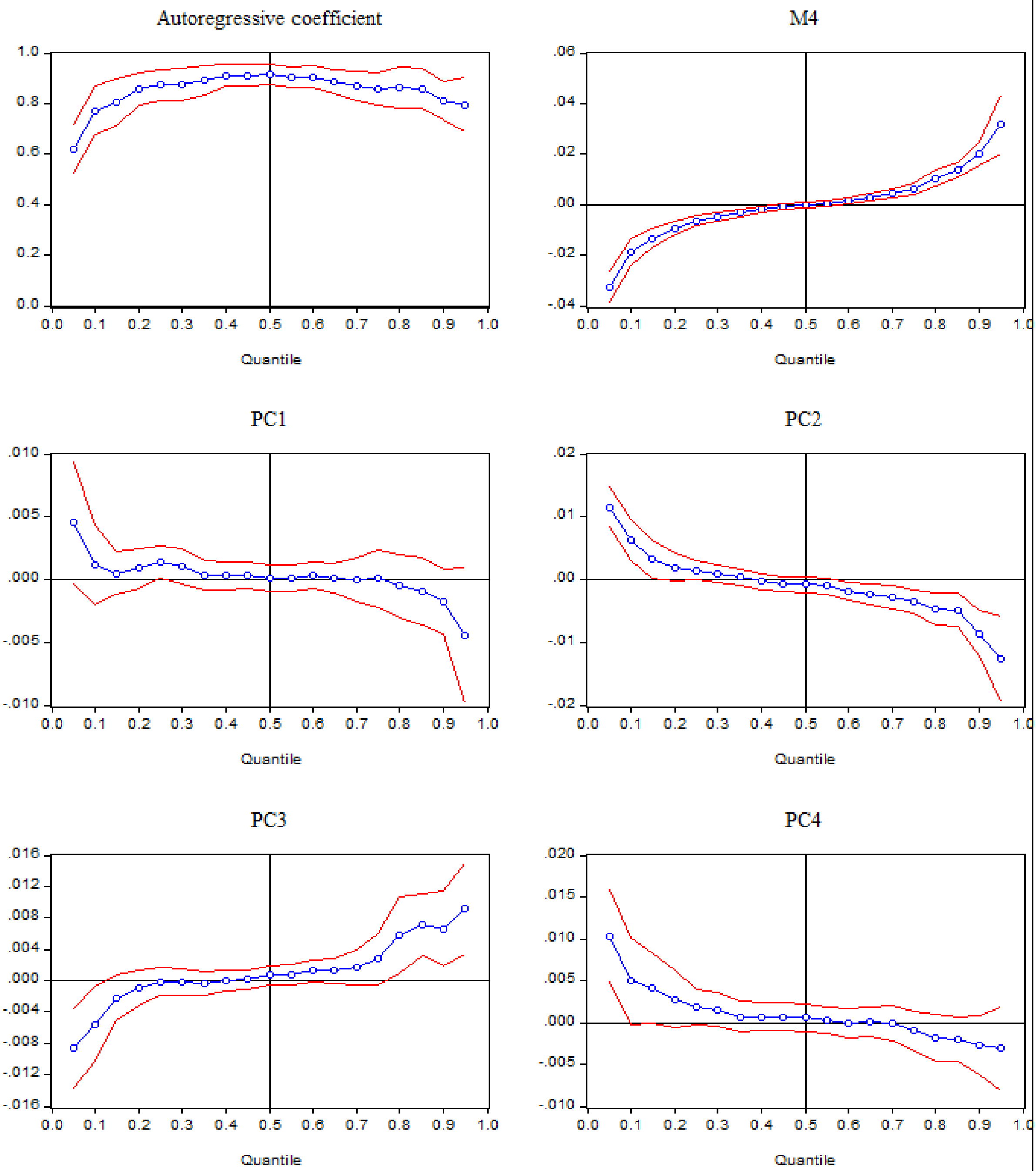


Figure 2.22: Quantile estimates from the multivariate model with principal components (6-month MA of M0)

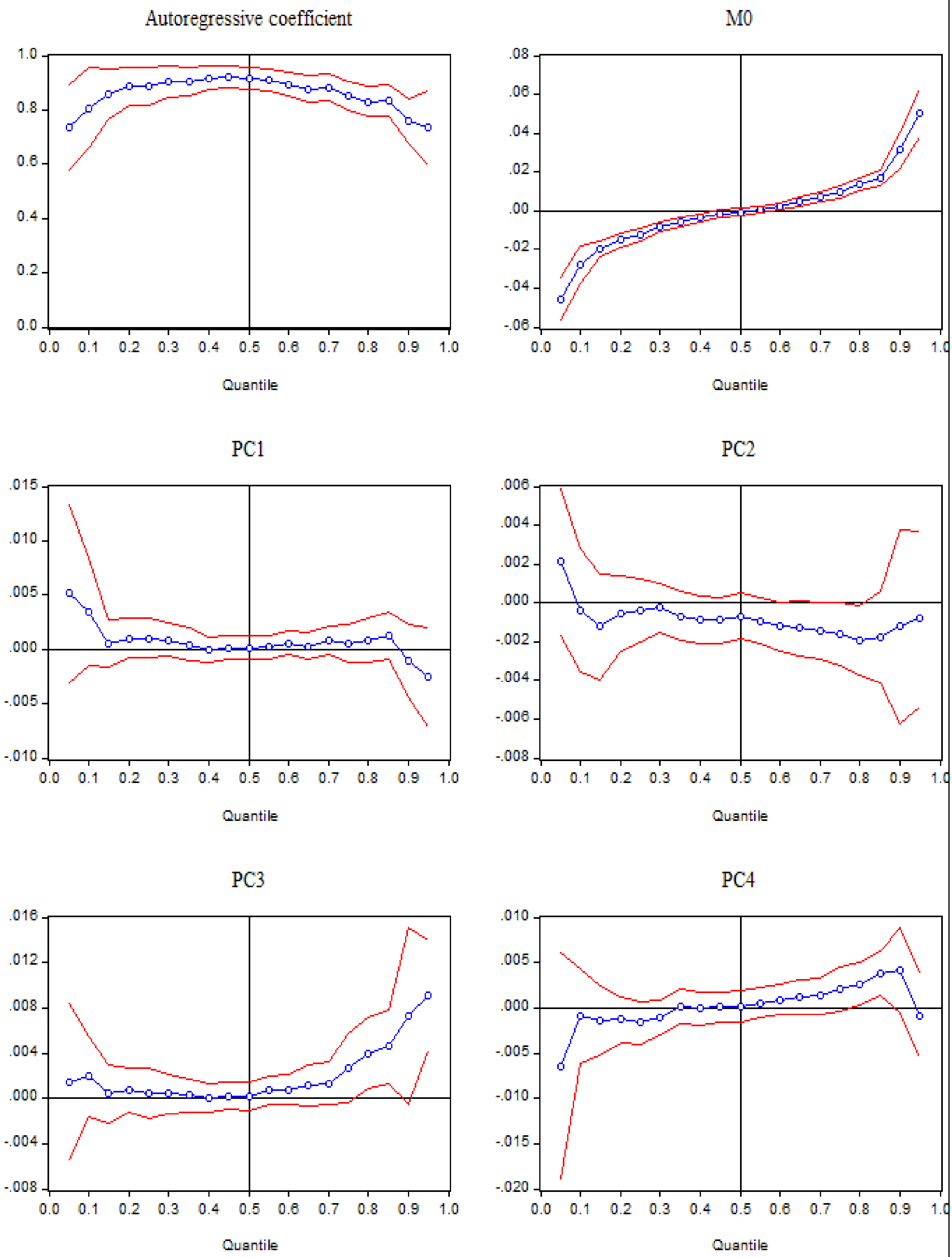


Figure 2.23: Quantile estimates from the multivariate model with principal components (6-month MA of IM)

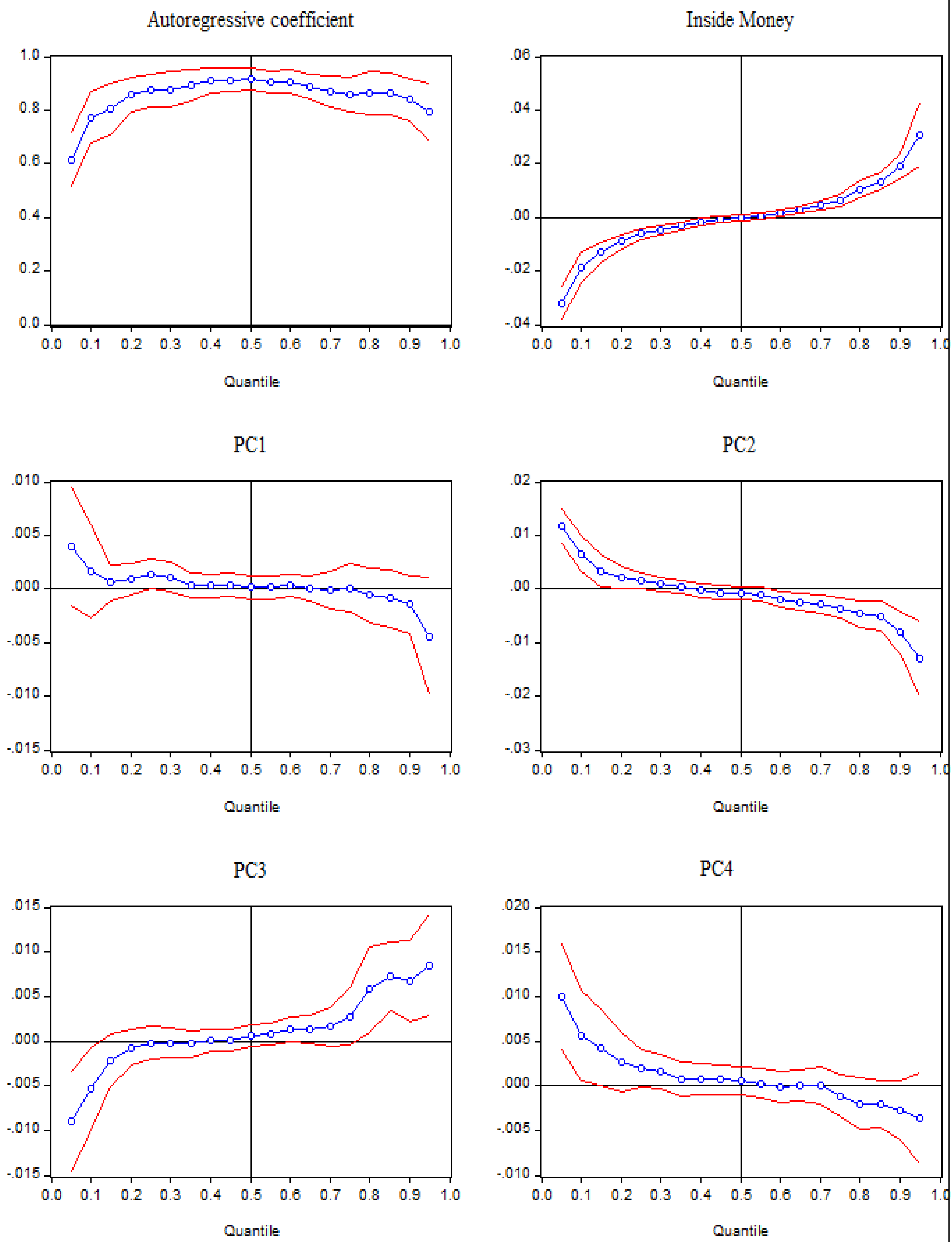
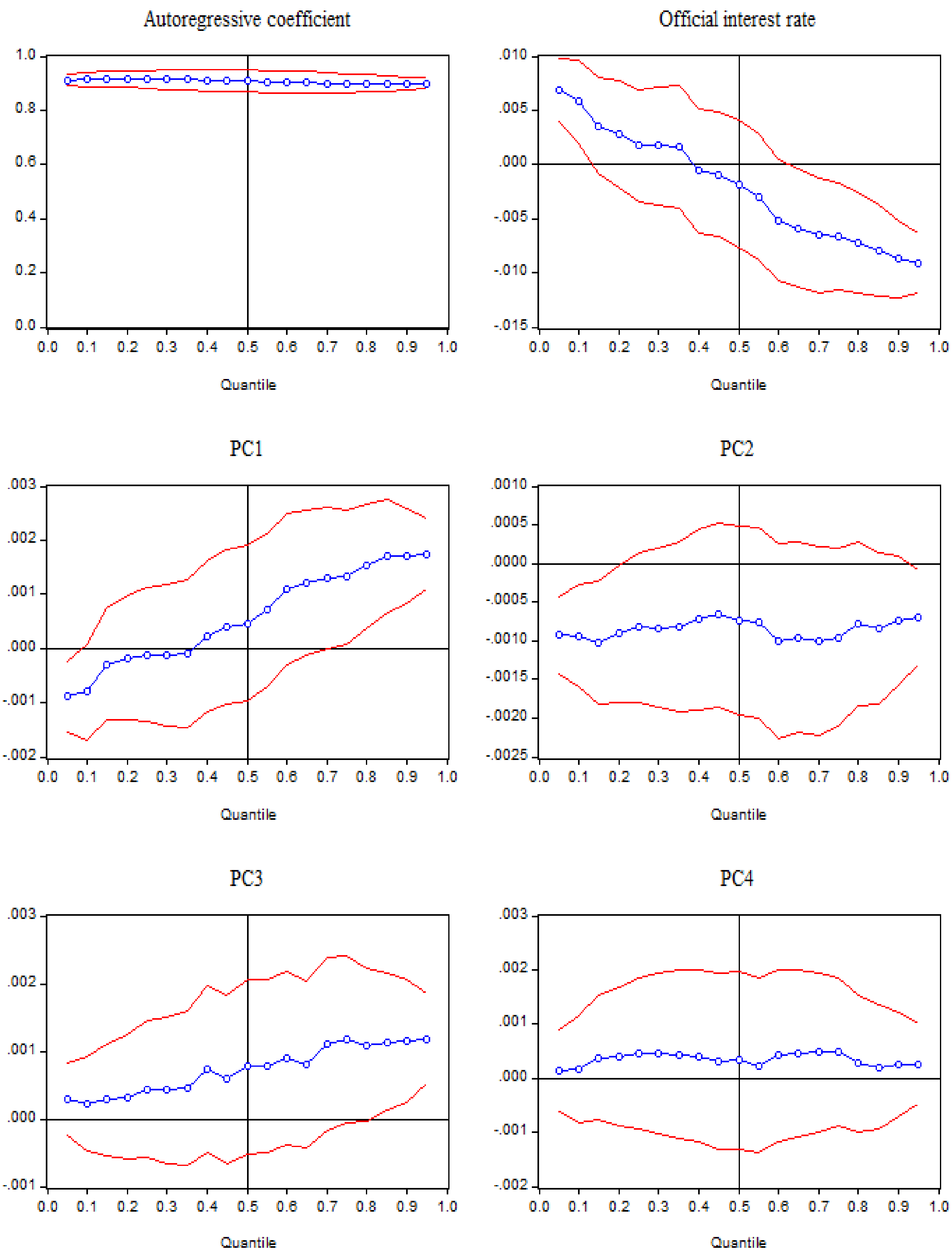


Figure 2.24: Quantile estimates from the multivariate model with principal components (6-month MA of official interest rate)



The figures show that monetary measures have predictive content for understanding changes in the variability of correlations, given the information content of other predictors and variability's own past history. The signs of coefficients at lower and higher percentiles are the same as the signs in the bivariate model. The results for the additional indicators which are worth mentioning are the following: First, the second principal component, which depicts economic recovery, has a negative impact on upper percentiles in all estimated regressions, suggesting that an economic recovery reduces the variability of dynamic correlations. Second, the third principal component, which reflects the inflation effect, has a positive impact on the upper percentiles in regressions with monetary aggregates, suggesting that an increase in inflation increases the variability of dynamic correlations.

### *Bivariate regression results with endogenous monetary variables*

We have also examined the contemporaneous effect of monetary measures on the variability of dynamic correlations using Amemiya's (1982) two-stage LAD (2SLAD) estimator. Initially, we estimate with LAD an auxiliary regression of the form: $x_t = \theta_0 + \theta_1 x_{t-1} + \theta_2 \tilde{\rho}_{t-1} + \zeta_t$, where $x_t$ denotes the growth rates of M4, M0, IM, and the change in the official rate, and get the fitted values $\hat{x}_t$. Then, we insert these values in quantile equation: $Q_{\tilde{\rho}_t}\left(\tau | \tilde{\rho}_{t-1}, \hat{x}_t\right) = \left(1, \tilde{\rho}_{t-1}, \hat{x}_t\right)' \vartheta(\tau)$, which is also estimated with LAD (Wooldridge, 2007). Apart from estimating the above auxiliary regression, we have also estimated another version which uses three-month and six-month moving averages of monetary regressors. Table 2.10 reports estimates from bivariate quantile regressions using fitted values of monetary variables obtained from the above auxiliary regression.[8]

| **Table 2.10: Quantile two-stage estimates from the bivariate model** | | | | | |
|---|---|---|---|---|---|
| | *Quantiles* | | | | |
| Regressors | 0.05 | 0.25 | 0.5 | 0.75 | 0.95 |
| $\tilde{\rho}_{t-1}$ | 0.635***<br>(0.076) | 0.862***<br>(0.035) | 0.915***<br>(0.021) | 0.898***<br>(0.025) | 0.794***<br>(0.028) |
| $\Delta\hat{m}4_t$ | -0.029***<br>(0.005) | -0.007***<br>(0.001) | -0.001***<br>(0.001) | 0.005***<br>(0.001) | 0.026***<br>(0.002) |
| | | | | | |
| $\tilde{\rho}_{t-1}$ | 0.712***<br>(0.047) | 0.865***<br>(0.031) | 0.918***<br>(0.021) | 0.882***<br>(0.026) | 0.665***<br>(0.045) |
| $\Delta\hat{m}0_t$ | -0.048***<br>(0.004) | -0.012***<br>(0.002) | -0.001<br>(0.001) | 0.008***<br>(0.001) | 0.050***<br>(0.006) |
| | | | | | |
| $\tilde{\rho}_{t-1}$ | 0.631***<br>(0.077) | 0.865***<br>(0.035) | 0.914***<br>(0.021) | 0.900***<br>(0.025) | 0.809***<br>(0.028) |
| $\Delta i\hat{m}_t$ | -0.028***<br>(0.005) | -0.007***<br>(0.001) | -0.001<br>(0.001) | 0.005***<br>(0.001) | 0.025***<br>(0.002) |
| | | | | | |
| $\tilde{\rho}_{t-1}$ | 0.917***<br>(0.009) | 0.916***<br>(0.017) | 0.907***<br>(0.021) | 0.896***<br>(0.019) | 0.895***<br>(0.010) |
| $\Delta\hat{R}_t$ | 0.004**<br>(0.002) | 0.001<br>(0.002) | -0.001<br>(0.002) | -0.004**<br>(0.002) | -0.008***<br>(0.002) |
| **Notes:** Sample size: 384. Numbers in parentheses are SEs. ***, ** indicate significance at 1% and 5% level respectively. | | | | | |

We observe that monetary regressors have a statistically significant contemporaneous impact on the variability of dynamic correlations in lower and upper quantiles. The effects of M4, M0 and inside money are negative in lower quantiles and positive in higher quantiles, while the effect of the official interest rate regressor is positive in lower quantiles and negative in higher quantiles. These findings are qualitatively the same with those obtained from regressions using moving averages of monetary variables.

[8] When in the quantile regression were used the fitted values from auxiliary regressions with three-month and six-month moving averages of monetary variables, the results were qualitative the same as those reported in Table 2.10.

## 2.6. Conclusion

Several empirical studies have examined the monetary policy effects on individual asset returns, and the influence of macroeconomic and risk factors on stock-house correlations. This chapter adds to this literature by examining the role of monetary policy variables in the UK in predicting changes in the dynamic covariance between stock and house returns. The importance of this topic stems from the fact that firstly stock and house markets constitute transmission mechanisms of monetary policy, and secondly stocks and houses constitute large components of the total wealth of UK households, thus affecting consumption and investment decisions.

Initially, we use the DCC model to filter the dynamic covariance between asset returns. In turn, we use a least squares with breaks method to filter the variability of dynamic conditional covariance between the stock market return and house market return. The variability reflects periods of positive and negative deviations from trend. If deviations are large, boom and bust outcomes are present and imply risk. Managing risk means having information about the whole distribution of these possible outcomes. To get information about these outcomes, we relate changes in dynamic covariance with predetermined developments in monetary variables. In particular, we use a quantile approach to examine whether monetary variables contain information for forecasting changes in the dynamic covariance of stock-house returns across the entire conditional distribution of the response variable. The following results emerge from the first part of the empirical analysis. First, the DCC model is a mean-reversion process which exhibits a high degree of persistence. Secondly, the dynamic correlations of stock-house returns have displayed a different profile during the sample period. Initially, a sharp decline occurred at November 1987 when the stock price index in the UK attained its lowest level as a result of the world stock market crash, indicating that stock and house returns moved opposite. Since the beginning of 2007, the correlations became positive, even during the global financial crisis 2007-2009, indicating that the stock return and house market return have moved in the same direction. From the second half of 2010, the upward trend has been reverted and the two asset returns exhibit a tendency to return back to its mean value by the end of the sample.

The second part of the empirical analysis examines whether fluctuations in stock-house returns are driven by developments in monetary variables, given the contribution of other macroeconomic variables. The following results emerge from the second part of the empirical analysis. First, the dynamics of the variability of dynamic correlations exhibit an asymmetry in the persistence profile. In particularly, in most cases the value of the autoregressive coefficient increases as we move from lower quantiles to median where it achieves its highest value, and then it falls as we move from median to higher quantiles. Thus, positive and negative shocks have less persistent effects on the variability of dynamic correlations in the lower and upper quantiles. Second, the three-month and six-month moving averages of money measures have a statistically significant impact on current variability of dynamic correlations, and the sign of their impact is positive in the higher quantiles and negative in the lower quantiles. This asymmetric effect implies that when changes in dynamic correlations are positive, an increase in the money supply and its inside component will increase the variability of correlations of stock-house returns, thus increasing the possibility of a boom episode. On the other hand, when changes in dynamic correlations are negative, an increase in the money supply and its inside component will decrease the variability of stock and house returns, thus increasing the possibility of a bust episode. This finding shows that movements in M4, M0 and inside money are important to understanding future movements in the variability of dynamic correlations between the stock market return and house market return, given the information provided by variability's own past history. Third, the three-month and six-month moving averages of changes in the official rate have a statistically significant impact on changes in dynamic correlations, and the sign of its impact is negative in the higher quantiles and positive in the lower quantiles. This asymmetric interest rate effect implies that when variability is positive, a fall in the official rate will further increase the variability of stock-house returns, thus increasing the possibility of a boom episode. On the other hand, when variability is negative, a fall in the interest rate will decrease the variability of stock-house returns, thus increasing the possibility of a bust episode. This finding shows that movements in the official interest rate have information content that is useful to understand future movements in the variability of dynamic correlations between the stock market return and the house market return, given the information provided by the variability's own past history.

We have also examined the predictive content of monetary measures in the context of multivariate equations which include six additional indicators of original data and indicators extracted as principal components from these six control variables. The results have indicated that all monetary measures have predictive content for understanding changes in dynamic correlations, given the information content of other predictors and variability's own past history. The signs of coefficients at lower and higher percentiles are the same as the sign from the bivariate model. Finally, the results for the additional indicators derived from PCA suggest that an economic recovery reduces the variability of dynamic correlations in all estimated regressions, while an increase in inflation increases the variability of dynamic correlations in the upper percentiles in regressions with monetary aggregates.

In a nutshell, the conclusion which results from quantile analysis and answers the research question is that developments in monetary policy measures in the UK, reflected in changes in monetary aggregates M0 and M4, and the official interest rate, have characterized the stylized facts of dynamic covariance between stock and house returns during the sample period, and hence provide valuable predictive content for the variability of the dynamic covariance. This evidence from the UK adds to the existing literature that stock and house markets constitute mechanisms though which monetary policy decisions affect the macroeconomic environment, and thus these facts should be taken into consideration when policy decisions are designed.

# Chapter 3

# Real output growth and public debt for three centuries: A Markov switching analysis

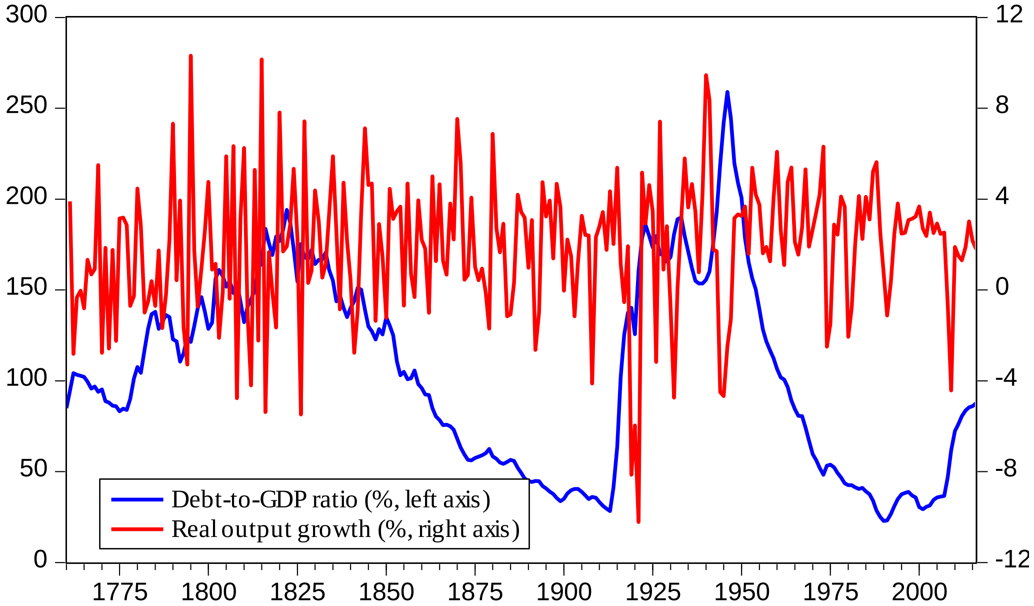
## 3.1. Introduction

The global financial crisis of 2007-2009 has produced disappointing economic outcomes which have affected the real economy (Bernanke, 2018). The transformation of the financial crisis to sovereign debt crisis has questioned the long-term fiscal sustainability in advanced and developing economies, given the fact that the debt-to-GDP ratios have reached high levels. A crisis might bring negative effect into the real economy not only by destroying financial wealth and obstructing financial intermediation (Reinhart and Reinhart, 2010), but also by aggravating fiscal positions that brings public debt beyond certain levels which may not be sustainable (ECB, 2015). In the aftermath of the financial crisis, the importance of fiscal sustainability for long-term economic performance has gained a new research interest, and thus several studies have examined the relationship between public debt and the real economy. However, the empirical literature has examined the link between public debt and real output growth without capturing an empirical fact that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods. This empirical regularity cannot be explained by linear models used in the literature which cannot capture these distinct facts, but instead a nonlinear model capturing these distinct patterns in the data is regarded more appropriate. A nonlinear model which is suitable to analyze different dynamic structures of the data over a time period and allows switching between these structures is the notorious Markov switching model (Hamilton, 1994, 2010). This chapter adds to literature by fitting a Markov switching model to annual real output growth rate from the middle of eighteenth century until 2016 in order to examine changes in mean during a period of 257 years. The overall objective is twofold: Firstly, to identify changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. Secondly, to examine whether the growth rate of the debt-to-GDP ratio is a leading indicator of the time-varying transition probabilities between expansions and contractions of the real output growth.

The importance of this research stems from the fact that the link between debt and output growth is an important topic from a policy perspective, as the establishment that increases in the growth rate of the debt-to-GDP ratio lead to switching from expansion to contraction regimes will support the arguments for long-run fiscal consolidation.

Figure 3.1 plots the growth rates of the real output and the debt-to-GDP ratio. From 1760 to 2016, the debt-to-GDP ratio in the UK has fluctuated widely. From 96% in 1860, it was reduced to 23% in 1913. Then, it increased to 259% in 1946, dropping thereafter to 23% in 1991, and increased again to about 88% in 2016. During the same period, the annual growth rate of the UK economy, as measured by the real GDP, has also fluctuated widely in a range of +/-10%. The standard deviation of the growth rate of the debt-ratio and the growth rate of real GDP is equal to 8.14 and 6.17 respectively. The correlation coefficient between the growth rates of the debt ratio and real GDP equals -0.39 ($t$-statistic=-6.8, $p$-value=0.00), showing that the government debt and real output have a contemporaneous moderate negative association.

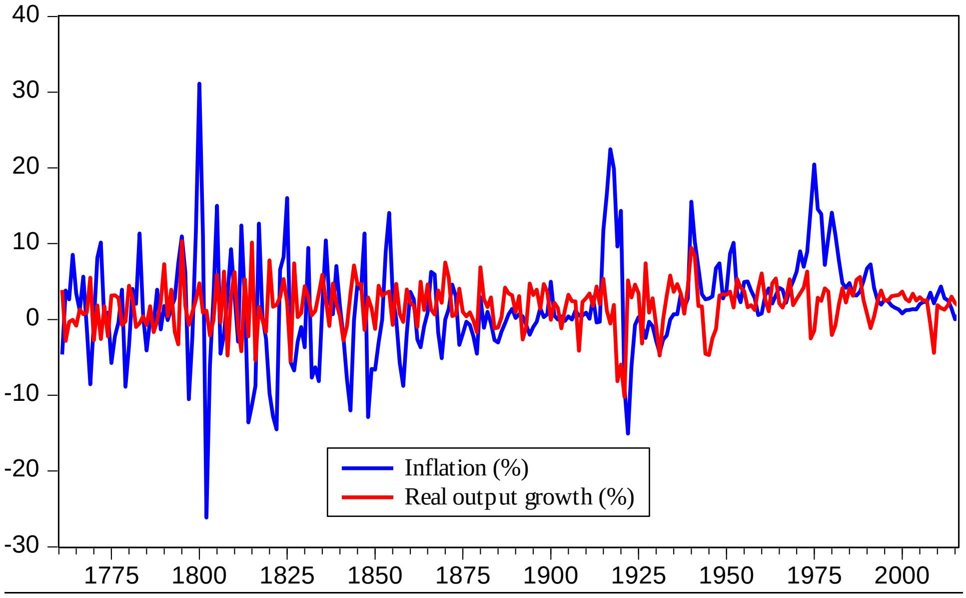
Figure 3.1: The debt-to-GDP ratio and the growth rate of real output




In the literature review, we present a number of papers which have established that high and variable rates of inflation are also detrimental to real output growth. During the last three centuries, the inflation rate of the UK economy has fluctuated widely, as it is evident from Figure 3.2 which plots the inflation rate and the growth rate of real GDP. From 31% in 1800, it has been thereafter significantly reduced to a negative rate in 1914. Then, it increased again to 22% in 1917, dropping thereafter gradually to almost 0% in 2016. Given the fact that the UK inflation rate has been very volatile during the sample period, we have included inflation in our model as an additional determinant of real output growth.

The rest of the chapter is organized as follows. Section 3.2 presents the literature review. Section 3.3 discusses the theoretical issues pertaining to the research question. Section 3.4 presents the methodology. Section 3.5 shows the empirical analysis. Section 3.6 concludes.

Figure 3.2: Inflation rate and growth rate of real output




## 3.2. Literature Review

This section discusses the empirical literature which examines the relationship between public debt and economic growth on the one hand. The main conclusion which results from the relevant literature is that the empirical studies have examined the link between public debt and real output growth without capturing an empirical fact that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods. This empirical regularity cannot be explained by linear models used in the literature which cannot capture these distinct facts, but instead a nonlinear model capturing these distinct patterns in the data is regarded more appropriate. This discussion reveals the gap in the literature referring to a non-linear analysis of the asymmetric behavior over real output expansions and contractions, and the role of the debt-to-GDP ratio in driving the transition probabilities between these dynamic structures, and this chapter aspires to fill this gap. In the literature review, we also present a number of papers which have established that high and variable rates of inflation are also detrimental to real output growth. Given the fact that the UK inflation rate has been very volatile during the sample period, we have included inflation in our model as an additional determinant of real output growth.

We will begin the overview with a paper which examined the effects of fiscal policy on the U.K. economy for the period between the beginning of the eighteenth century and the World War I. In particular, Barro (1987) investigated how changes in government expenditure affected a number of economic variables, such as money quantity, interest rates, level of prices and budget deficits. He firstly found out that rise in this expenditure led to an increase in interest rates, while for the money quantity, the effect was also positive but only during the periods of gold standard's abolition. In a similar note, inflation was also affected positively by the government expenditure during the above periods. Regarding the budget deficit, this one was accumulated during the whole period in general. However, during some specific periods when there were compensations which were paid to slave-owners and when there was fight among the politicians over the income tax, the budget deficit was totally separated from the effects of the government spending. Reinhart and Rogoff (2010) examined for the existence of a relationship between debt, growth and inflation. They used data from 44 countries over a sample period of 200 years around. They provided evidence by splitting the countries into two separate groups, the advanced one which included 20 countries and the emerging market one which contained 24 countries. They also decided to divide the debt into different categories, such as periods when there was low, medium, high and very high level debts in order to examine for each case how the debt level affects the economic growth. They have found initially that for the advanced countries' group there is no clear relation between debt and growth when the debt level is low, medium and high. Once the debt reaches and surpasses the 90% of GDP level threshold, it was observed that the median of economic growth declined by 1% less than the other debt level categories. In the group of the emerging markets countries, the median growth was found to be about 4 to 4.5% when the debt level was low, medium or high and once it was very high, the median growth was reduced to 2.9%. As a result, they observed similarities to both groups of countries and their main finding was that very high debt level is linked to very low economic growth. Kumar and Woo (2010) studied the effect of high pubic debt on the economic growth in the long-run. For their research, they used data for 38 developed and emerging countries for a period of 38 years and the main data source was the Penn World Table. They started their estimation with a multi-country OLS regression, where they included the logarithm of real per capita GDP as the dependent variable while the independent ones were a group of economic and financial variables and also the initial government debt. They also proceeded to investigate if the results they have found were robust by including more macroeconomic variables. Their main result was that there is an inverse link between debt and growth. More specifically, if there is a 10% rise on the debt-to-GDP ratio, then the annual real per capital GDP will decrease by about 0.2%. They have also found that when the debt level is very high, this has severe negative effect on the economic growth, which is also shown in the reduction of the labor productivity growth. Pattillo, Poirson and Ricci (2011) examined the effect of external debt and debt reduction on economic growth. For this occasion, they focused on panel data from 93 developing countries during the sample period of 1969-1998 and they used several econometric methods, such as fixed effects and the Generalised Method of Moments (GMM), in order to check how the results were affected by endogeneity and dynamic panel biases. They initially found out that there is indeed a nonlinear link between debt and growth, which seems to be negative in average when the debt level exceeds the 160-170% of exports threshold and 35-40% of GDP one. Regarding the marginal impact, instead of the average one, the thresholds turned to be half of the above ones. All in all, it was difficult for the exact threshold to be calculated, because of various limitations in the dataset used. In addition, they proved that if a country has a debt level which is much higher than the ultimate threshold, then having a doubled debt could lead to a reduction of the economic growth by half percent or less. Finally, they included the investment to check if it plays any role on the effect of debt on growth and the results suggested that this role is very insignificant, implying that investment is not the transition mechanism through which external debt declines economic growth. Delong and Summers (2012) investigated the effects of zero bound interest rate fiscal policy on an economically depressed country, such as the U.S.A., which means that there is high unemployment and low levels of output compared to the potential ones. Their main findings were that this kind of policy raised the Keynesian multiplier significantly higher than normal. Furthermore, in case the government had more spending, this would likely increase the future output, while austerity policies would harm the economy in the long-run. Baum, Checherita-Westpal and Rother (2012) surveyed the link between public debt and the growth of per-capital GDP in 12 euro area countries over a period of about 40 years. Their main data source was the European Commission and their main control variables were initial level of per capita income, the ratio of investment over saving-to-GDP and the growth rate of population. They used a quadratic equation in order to determine whether the examined link was linear or non-linear. Their main findings were that the relationship between the 2 variables was found to be non-linear and the turning point after which, the effect of debt on growth became negative was on average around 90-100% of GDP, however according to the confidence intervals, this level could go as low as 70%. This is an indication that the current debt levels of many of the examined countries could prove to be already harmful for their national economic growths. Finally, the variables through which the link proved to be non-linear were total factor productivity, public investment and private saving. Reinhart, Reinhart and Rogoff (2012) focused on how periods of high levels of debt affected growth and interest rates. For this purpose, they specifically studied 26 events in a group of advanced economies during the period of 1800-2011 in which at least 90% debt to GDP was observed for at least 5 consecutive years. They observed that during the majority of these events, the economic growth was found to be significantly 1.2% lower than during periods where the debt to GDP ratio was below 90%. That also gave evidence that countries with prolonged slow economic growth and high debt would find it more difficult to run away from their debt obstacles. Furthermore, if a country suffered from excessive levels of debt for 5 years, it is high likely to suffer for 10 more years at least. It was also found that interest rates were significantly raised because of the debt overhang crisis. Panizza and Presbitero (2013) examined whether the government debt has indeed a negative effect on economic growth while it reaches high levels in developed countries which was suggested by a number of theoretical and empirical papers. They found that a causal relationship between the two variables was not established and also the debt level of 90% which was highly used by a lot of papers as the threshold beyond which the debt has significantly negative impact on growth was not vigorous. They also addressed the need the debt to be defined correctly, for example whether should be the gross or net one to be used in the context. They suggested that in order to prove that there exists a strong link between debt and growth, the analysis should be based on cross-country heterogeneity analysis and should also focus on the paths and instruments via which the debt affects the economic output. Ghosh et al. (2013) investigated how high the debt can increase without the intervention of any fiscal policies. For this purpose, they have created a new variable called "fiscal space", which is the difference between the current debt level of the examined country and its debt limit, in which the fiscal policies cannot control the debt anymore. They extracted data from 23 developed countries for a period of 38 years. They firstly found that the debt has a non-linear positive effect on the primary balance when the debt level was average, but it turned to be negative, once the debt level reached high levels. In addition, they have found that the debt limits were quite different across the estimated countries and also the fiscal space analysis indicated that for the European countries which have suffered the most during the past years plus Japan, fiscal manipulation could not be implemented, while it was possible for the Scandinavian counties, Australia and Korea. Therefore, with their research, they tried to provide evidence about which countries have available fiscal space and in case this one is limited how it can be extended by the government or other institutions. Eberhardt and Presbitero (2013) investigated the link between government debt and growth in the long-run for numerous countries. Their principal variables were total public debt stock, capital stock and GDP and the main data source was the World Bank World Development Indicators. For the empirical analysis, they used linear dynamic and asymmetric dynamic models in order to confirm the existence of a long-run relationship and also to show the causation direction of this relationship. They also used techniques to determine the heterogeneity among the countries. Their main findings were that there were significant differences between debt and growth among the examined countries, however they did not find any proof about nonlinearities inside the countries for all of them. More specifically, the higher the debt problems a country had the lower the debt coefficient appeared to be in the long-run. They also found that the general examined relationship seemed to be different across all the examined countries, which means that countries should follow different policies to combat their debt issues. Finally, they observed that the 90% debt level which was widely used by other authors seemed not to be valid for their research, as the coefficient of the debt change could take positive and negative values at the same time. Kourtellos, Stengos and Tan (2013) examined whether there is any linear or non-linear link between debt and growth from the viewpoint of specific political regimes. They used 3 decades data for a sample of 82 countries between 1980 and 2009. Moreover, their dependent variable was the growth rate of real per capita GDP, while the independent ones were, except for the debt-to-GDP ratio, the trade openness, the inflation and a measure for the government size. Their empirical methodology was a structural threshold regression model based on the Solow growth one. Their results suggested that there is weak observation of a nonlinear link between growth and debt. On the other hand, they proved that the existence of particular regimes can be responsible for the relationship between the two variables. In particular, if the political regime is not very democratic, then high levels of debt can lead to very low growth, however in the case of democratic regimes, the effect is more neutral. Pescatori, Sandri and Simon (2014) investigated if there is any government debt threshold above which the growth could be weakened. They based their research on previous literature which was divided into two groups. The first one suggested that when the debt reaches high levels, this has a negative effect into growth; however the second one opposed this theory by implying that high debt levels and negative growth are not directly related. Thus, the authors contributed to this debate by using a totally new methodology which took into account longer periods of data, which was dated from 1875 regarding 19 advanced countries, in order to examine the relation not only in the short term, but also in the medium term. They extracted their data from IMF Fiscal Affairs Department. They found out that there is no clear link between debt and growth once debt reaches high levels, especially in the medium run. They also proved that the existence of the debt trajectory is more significant that the debt level itself so as to understand the growth movements. However, they noted that high debt levels are responsible for making the debt volatile and as a result debt should always be taken into account, no matter how strong or weak is the link between the two variables. Égert (2012) contributed to the existing literature about the negative link between debt and economic growth by putting into test the dataset of Reinhart-Rogoff (2010) and checking if there is indeed a threshold after which the effect of debt on growth is very significant. For this case, he extracted data for 29 OECD countries which were spanned between 1960 and 2010. Firstly, by applying a non-linear threshold model, he found out that the existence of a negative link between debt and growth is not completely valid. More specifically, there were significant variations of the nature of the link between the two variables and also the negative link was found to be much lower than the widely proposed 90% threshold. As a result, he suggested that the latter threshold after which the debt largely affects negatively the growth should not always be taken into account as it could be for example even lower than 60%. Furthermore, the nonlinear link can also be different across various samples, time periods or countries.

We now turn to the empirical literature which examines the relationship between inflation and economic growth. Roubini and Sala-i-Martin (1992) surveyed how a financial repression can affect economic growth. The intuition behind this policy was that, according to some governments, financial repression may lead to a rise for money demand which in turn may lead to rise in revenues derived from inflation. They mainly focused on the Latin American countries, as this region seemed not to grow as much as the rest of the planet. To begin with, they did their empirical work on 98 countries during 1960-1985 and they used a lot of independent variables such as, the GDP value in 1960, the human capital value in 1960, the government expenditure, the deformation in the prices of investment goods as well as political and social ones. The empirical findings showed that almost all of the above regressors affected the economic growth in a negative way. They later focused on the financial repression policies especially for the Latin American area. Indeed, during the above period of 1960-1985, the Latin American countries appeared to have slower economic progress due to the existence of financial repression or underdevelopment. Bullard and Keating (1995) studied the relationship between inflation and growth for lots of countries after the Second World War. Specifically, they used data for 58 countries which were obtained from the International Financial Statistics. The main variables included into their empirical model were the real gross domestic product and the gross domestic product deflator. The next step was to divide this amount of countries into smaller groups based firstly on the quality of the data and secondly on the availability of the variables concerned. In addition, the data period was roughly between 1960 and 1992. After applying permanent shocks to the inflation, they found out that in general these are not related to permanent changes into the output as well, however there were signs of some positive correlations between the two examined variables. The latter finding led to another conclusion that there was no valid long-run relationship which included permanent shocks to both variables at the same time. Chari et al. (1996) presented a summary of previous work about the effects of monetary policy instruments on growth from the empirical viewpoint and discussed about the use of quantitative models in which the growth rates are influenced. According to previous empirical work which they were based on, when the inflation increases by 10% in average, this will decrease the growth rate by about 0.2-0.7%. However, when they applied their own models, they found out that their findings were completely different. As a result, they had to change their focus on the monetary policy instruments by including money supply changes and financial regulations. With the latest additions, they found out that these regulations, especially when they interact with the inflation play a significant and negative role on growth, which means that policymakers should focus on the section of financial regulations rather than money issuing. Bruno and Easterly (1998) examined the relationship between high levels of inflation and economic growth and how robust is this link. They based their research on the fact that lots of papers stated that the relationship is negative, but they found out that this observation is not valid for all levels of inflation plus the data used must be of a high frequency. For this reason, they considered all levels of inflation which were at least 40% of annual rate. They also used data for 31 countries over a period of 1961-1994, where they noted of 41 high inflation events. Their main finding was that during periods of high inflation, the link between inflation and growth appeared to be significantly strong and negative and when inflation started to be reduced, then that was followed by higher economic growth. However, it could not be proved what the causal relationship was in the long-term between the two variables. Ghosh and Phillips (1998) investigated whether the negative relationship between inflation and economic growth could be observed not only in high inflation levels but also in moderate as well as low levels. For this purpose, they extracted data from 145 countries for the period between 1960 and 1996. They firstly found that when the inflation reached very low levels of about 5% and below, the relationship between inflation and growth turned to be non-linear and positive, whereas for moderate levels, such as those between 5 and 30%, is once again non-linear, significant and negative. However, they suggested that they exact turning point where the low inflation can have a positive effect on growth was difficult to be determined. Judson and Orphanides (1999) investigated the nature of the link between inflation and income growth as well as inflation volatility and income growth. Their analysis was based on a cross-country model which could incorporate 142 countries maximum; however for the majority of their regressions, they examined 119 countries. The duration of the dataset was 34 years covering the period between 1959 and 1992. Their contribution was firstly the inclusion of time series analysis and cross-section one at the same time and also the introduction of inflation volatility by using more frequent data. Regarding the inflation volatility, their empirical results showed that there was a strong and negative correlation between volatility and income growth for different levels of inflation and regarding the inflation itself, it is also linked to the growth in a negative way, but only for levels of 10% and above. One final remark was that they found out that inflation and inflation volatility affect the income growth in an independent way. Thus, their evidences support what the monetary policymakers suggest, which is negative influence of high inflation and volatile inflation on economic output. Rousseau and Wachtel (2001) examined the link between inflation and economic growth using data from 84 countries for the sampling period 1960-1995. As they wanted to examine the effects in the long-term, the frequency of this dataset was five-year averaging. Their empirical findings showed that with the introduction of the financial sector, the results are significant. More specifically, the financial development affects notably the economic growth, especially when the inflation level is medium. In respect of the direct link between inflation and growth, this is strong when the inflation level is high. Finally, the significant effect of the financial development on growth is not changed when the inflation is present at the same time. Khan and Senhadji (2001) studied what are the threshold effects on the link between inflation and growth by implementing nonlinear least squares method, since the thresholds are not known. Their dataset dated between 1960 and 1998 and it included 140 developing as well as industrialized countries. Their results suggested that there is indeed a threshold above which the inflation has a significant and negative effect on economic growth but the level of this threshold turned to be different for the 2 groups of the examined countries. For the developing countries, the level was 11-12%, while for the industrialized ones it was only 1-3%. Furthermore, this negative relation was found to be stronger when the data frequency was higher. As a result, this paper supported the idea that keeping inflation at very low rates could be beneficial for the economic output. Gylfason and Herbertsson (2001) investigated whether there was positive or negative effect of inflation on economic growth as it was largely argued by various economists. For this purpose, they used data for 170 countries from the World Data Bank covering the period 1960-1992, and for 145 countries from the Penn World Table. They observed that as the inflation increased, this had a detrimental effect on growth. More specifically, when this increase was from 5 to 50%, then the per capita GDP growth rate was lowered by 0.6% for the World Data Bank sample and 1.3% for the Penn World Table one. The link between the two examined variables was also found to be nonlinear. Finally, they also noted the importance of studying the channels through which this negative relation exists and they are related to money and finance. Ericsson, Irons and Tryon (2001) surveyed whether the negative relation between inflation and output growth is valid for both developing and developed countries. Initially, they discovered that the result found for the African and Latin American countries was not applicable for the G-7 ones. In other words, it was proved that the effect of inflation on economic growth was positive, however not statistically significant at the same time. In addition, doing the regression for the OECD countries exclusively, they found out that no long-run link between inflation and growth seemed to appear at all. The reason for that was the use of cross-country regressions, which took into account the growth levels of the variables rather than the actual levels. As a result, they supported that the relationship between inflation and economic growth is different for each examined country and economic policies should not be generalized upon these findings. Gillman and Nakov (2004) examined the causal relationship between inflation and output growth, using a model with endogenous money supply growth. For their study, they used data from Hungary between 1987 and 2004 and Poland between 1986 and 2004. Their main methodology included the use of VAR models with the inclusion of structural breaks and also the Granger causality method in order to determine the direction of the causality between inflation and growth. They have found that money supply affected the economic growth with the presence of the inflation tax. About the Granger causality part, for both countries, it was observed that the direction of Granger causality is from money supply to inflation and from inflation to growth. More specifically, a rise in the money supply caused inflation to rise as well and this in turn caused economic growth to decrease. Furthermore, they found that in the case of Poland, there was also Granger causality from inflation and output growth towards money growth. Gillman and Kejak (2005) used a broad range of different models in order to examine the relationship between inflation and economic growth. They initially tried to collect all the different models with either physical or human into a common one which contained both. The purpose of this technique was to find out which variables are the most responsible for the determination of the link between inflation and economic growth. Another important point that was made was the determination of the role of inflation either as a tax on physical or human capital. This led to connect the inflation-growth relationship with the Tobin (1965) effect. They proved that for the physical capital tax models the Tobin effect was negative, whereas for the human capital tax ones it was positive and finally with the inclusion of both, the overall Tobin effect was again positive. It was also presented that there was a strong non-linear relationship between the two variables which was mainly because of the inclusion of the elasticity of money demand. Lopez-Villavicencio and Mignon (2011) studied the nature of the relationship between inflation and growth for countries with different levels of income, from high income OECD to emerging countries. More specifically, 44 countries were used in their study for the period between 1961 and 2007. Their empirical analysis was based on panel smooth transition regression model and also the dynamic GMM one for panel data. They initially observed the existence of a non-linear link between inflation and growth. The most important finding though was the calculation of the threshold above which inflation impacts negatively the economic growth. However, this threshold was found to be different across the various groups of countries. For example, for the high income one the threshold was proved to be 2.7%, while for the developing ones was 17.5%. On the other hand, for the first group, the positive effect of inflation on growth when it was below 2.7% seemed to be significant which did not happen with the second group for inflation levels below 17.5 %.

## 3.3. Theoritical Issues

There are many theoretical models which have shown that fiscal policy constitutes a determinant of real output growth. According to the conventional approach to fiscal policy discussed in Elmendorf and Mankiw (1999), a temporary tax reduction that creates a budget deficit, which is financed by issuing public debt, will increase disposable income and wealth, and consequently consumption expenditure. If nominal rigidities are present, such as sticky wages, sticky prices, or informational misperceptions, then the increase in aggregate demand will increase the demand for labour that will boost real output in the short-run. In order to explain the long-run effects of the debt policy, we will resort to the macroeconomic identity which implies that the national saving, consisted of the private and public savings, are equal to private investment and net capital outflows. If the fall in public saving that is occurred by the higher budget deficit would not be fully compensated by a rise in the private saving, which amounts to assume that the Ricardian Equivalence is not valid, then the national saving would fall, thus causing a fall in private investment and consequently a decline in the capital stock and economic growth, as proposed by the Solow model. The smaller capital stock will increase its marginal product and consequently its real return. In addition, the smaller capital stock would imply a lower labour productivity which would reduce the real wage. On the other hand, if the decline in national saving is matched by a fall in net capital outflows, the domestic currency would appreciate, thus resulting in a trade deficit which would reduce the real output in the long-run. It is worth noting that the high levels of public debt and the associated high level of interest rates may increase expectations of an accommodating monetary policy trying to reduce nominal interest rates. This policy would increase future inflation and nominal interest rates in the long-run, thus reducing real output. In addition, if people believe that the government would increase the future money supply in order to pay off in the future, then the resultant future inflation would cause inflation today which would be harmful for economic growth in the short-run (Cochrane, 2011a,b). In the context of the infinite horizon Ramsey-Cass-Koopmans model, a permanent increase in public expenditures reduces households' lifetime wealth and thus they do not adjust the time pattern of their consumption and consequently the saving rate. As a result, the capital stock remains constant and thus the long-run level of real output is not affected. On the other hand, a two-period version of the Diamond model, a permanent increase in public expenditures reduces the current consumption less than the increase in public spending, and as taxes are levied in the first period, households reduce their saving. As a result, the private investment is declined and consequently the capital stock and the long-run level of real output (Romer, 2012). Using an endogenous growth model, Greiner (2011, 2012) studies the role of debt policy and argues that the government debt has a positive impact on the real economy in the presence of wage rigidities and unemployment, if the debt is used to finance the private investment. In the absence of wage rigidities and an elastic labor supply, the government debt has a negative impact on private investment and economic growth.

## 3.4. Empirical Methodology

In this chapter, we use two empirical non-linear approaches to address two issues. Firstly, we estimate a Markov switching autoregressive model in order to identify changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. Secondly, we use an extended version of the above model by assuming that the probabilities between the two states are related to economic fundamentals in order to examine whether the growth rate of the debt-to-GDP ratio is a leading indicator of the time-varying transition probabilities between expansions and contractions. An alternative approach which assumes that the transition between the two states is smooth and depends on the past value of the dependent variable (Teräsvirta and Anderson, 1992; Teräsvirta, 1994) is not applied here because our research question is to examine whether lagged fundamenatsls have predictive content for a sharp regime switch between contractions and expansions.[9]

We start with the Markov switching model by assuming that the stochastic process of the growth rate of real output is modeled as an autoregressive first-order switching between expansions and contractions regimes:

$$y_t = c_{s,t} + \rho_s y_{t-1} + \sigma_s v_t \qquad (3.4.1)$$

where $y_t$ is the growth rate of real output, $c_{s,t}$ is the conditional mean of real output growth which is regime dependent with $s_t$ denoting a random variable which takes the value 1 during the expansion or high growth state and the value 2 during the contraction or low growth state, $v_t$ is *iid* standard normally distributed random error, and the standard deviation $\sigma$ may be regime dependent. The random variable $s_t$ is generated by a Markov stochastic process that is independent of all exogenous regressors and is defined by its transition probabilities:

$$p_{ij} = Pr\left(s_t = j \middle| s_{t-1} = i\right), \ \sum_{j=1}^{2} p_{ij} = 1, \forall i, j \in \{1,2\} \qquad (3.4.2)$$

The model is estimated by the maximum likelihood. The inference about the two values of $s_t$ takes the form of two probabilities and is performed iteratively. In particular, if $\lambda_{i,t-1}$ is the probability of being in state $i = 1,2$ at time *t*-1, given the set of observations $\Omega_{t-1}$ obtained as of time *t*-1 and the vector of parameters $\theta = \left(\sigma^2, \rho_1, \rho_2, c_1, c_2, p_{11}, p_{22}\right)'$, that is

[9] Deschamps (2008) compares the out-of-sample forecasting performance of Markov-switching and logistic smooth-transition autoregressive models for the US unemployment and concludes that both approaches provide very similar descriptions, but forecasting tests favor the smooth transition model. However, the ranking between these models depends on the transition delay parameter, and thus the two approaches can be complementary.

$\lambda_{i,t-1} = Pr(s_{t-1} = i|\Omega_{t-1};\theta)$, then $\lambda_{i,t-1}$ is used as input value in the next step of two probabilities at date *t*, given by: $\lambda_{j,t} = Pr(s_t = j|\Omega_t;\theta)$ for $j = 1,2$. In order to execute this iteration, we need the density function in each state:

$$\psi_{jt} = f\left(y_t|s_t = j,\Omega_{t-1};\theta\right) = \frac{1}{\sqrt{2\pi\sigma}} exp\left(-\frac{\left(y_t - c_j - \rho y_{t-1}\right)^2}{2\sigma^2}\right)$$

Thus, the conditional density at date *t* is: $f\left(y_t|\Omega_{t-1};\theta\right) = \sum_{i=1}^{2}\sum_{j=1}^{2} p_{ij}\lambda_{i,t-1}\psi_{jt}$, and consequently the two probabilities at date *t* are: $\lambda_{jt} = \sum_{i=1}^{2} p_{ij}\lambda_{i,t-1}c_{jt} \big/ f\left(y_t|\Omega_{t-1};\theta\right)$. This iterative procedure will evaluate the sample conditional log likelihood of the model: $\sum_{t=1}^{T} \log f\left(y_t|\Omega_{t-1};\theta\right)$ which is maximized to provide the value of the vector $\theta$.

In the above framework, the transition probabilities between the two states are not related to economic environment and consequently are considered to be fixed. Diebold et al. (1994) and Filardo (1994) have extended Hamilton's model by assuming that the probabilities between the two states are related to economic fundamentals. In this context, the probability of moving from one state to another is governed by the law:

$$p_{ij} = Pr(s_t = j|s_{t-1} = i, z_{t-1}, \theta),\ \sum_{j=1}^{2} p_{ij} = 1, \forall i, j \in \{1,2\} \qquad (3.4.3)$$

where $\theta = \left(\sigma^2, \rho_1, \rho_2, c_1, c_2, p_{11}, p_{22}\right)'$ is the vector of parameters, and $z_{t-1}$ is the economic fundamental observed at date *t*-1 and affect the transition probabilities.

In the empirical analysis, in order to answer the research question we proceed as follows: Firstly, we assume that the transition probabilities between the two states are fixed, in the sense that they do not depend on the exogenous regressors, and estimate the mean $c_{s,t}$, the propagation mechanism $\rho_s$ and the variance $\sigma_s$ of the real output growth rate in the two states. In this context, we examine whether the lagged value of the debt-GDP ratio has information content useful to predict movements in the growth rate of the real output in the two regimes. Secondly, we assume that the transition probabilities between the two states are time-varying, in the sense that they depend on the exogenous regressors, and estimate the mean $c_{s,t}$, the propagation mechanism $\rho_s$ and the variance $\sigma_s$ of the real output growth rate in the two states. In this context, we examine whether the lagged value of the debt-GDP ratio, which represents our economic fundamental $z_{t-1}$, has predictive content for the real output growth in moving from one state to another.

In oder to explore the main source of the nonlinearity, we estimate two nested variants of the model (3.4.1). The first one assumes that the mean and the dynamics of the real output growth change between the two states, and the second one assumes that the mean, the dynamics and the variance of the real output growth change between the two states.

## 3.5. Empirical Analysis

### 3.5.1. Data and unit root tests

The sample period spans from 1760 to 2016 for the following variables: real GDP, government debt to GDP ratio, the consumer price index, the consol (long-term bonds) yields and the real effective exchange rate. The growth rate of real GDP, $y_t$, is defined as the first-difference of the log of real GDP between year $t$ and year $t$-1. The growth rate of debt-to-GDP ratio, $d_t$, is defined as the first-difference of the log of the debt ratio between year $t$ and year $t$-1. The inflation rate, $\pi_t$, is defined as the first-difference of the log of the consumer price index between year $t$ and year $t$-1. The real long-run interest rate, $r_t$, is defined as the consol (long-term bonds) yields at year $t$ minus the inflation rate. The growth rate of real effective exchange rate, $e_t$, is defined as the first difference of the log of the real exchange rate between year $t$ and year $t$-1. All the data have been obtained from the Federal Reserve Economic Data St. Louis FRED, which have been constructed by the Bank of England.[10]

Initially, we test whether our series concerned are stationary processes using three unit root tests, namely Dickey and Fuller (1979), Philips and Perron (1998), and Kwiatkowski et al. (1992). The results presented in Table 3.1 indicate that the growth rate of real output, the growth rate of debt ratio, the inflation rate, the change in the real long-run interest rate, and the growth rate of the exchange rate are all stationary processes.

**Table 3.1: Unit root tests**

| Variables | Lags ADF | ADF | Lags PP | PP | Lags KPSS | KPSS | Result |
|---|---|---|---|---|---|---|---|
| $y_t$ | 0 | -14.776*** | 10 | -14.776*** | 8 | 0.3497 | **I(0)** |
| $d_t$ | 0 | -8.121*** | 10 | -7.907*** | 7 | 0.080 | **I(0)** |
| $\pi_t$ | 3 | -6.24*** | 13 | -9.609*** | 1 | 0.182 | **I(0)** |
| $\Delta r_t$ | 6 | -11.181*** | 34 | -46.588*** | 34 | 0.089 | **I(0)** |
| $e_t$ | 5 | -9.956*** | 24 | -15.856*** | 22 | 0.210 | **I(0)** |

**Notes:** Sample size: 257. For the ADF test, the lag selection of the dependent variable was based on SIC with a maximum lag=10. For the other two tests, the lags' column refers to the lag selection parameter, based on Newey-West automatic process using Bartlett Kernell spectral estimation method, which indicates how many autocovariances are used in taking the parametric estimators of the smoothness of the sprectral density. I(0) stands for integration of order zero (stationary) process. The critical values for the three tests at 1% significance level are as follow: Augmented DF (ADF) = -3.460, Phillips-Perron (PP) = -3.459, Kwiatkowski et al. (KPSS) = 0.739. *** indicate significance at 1%.

[10] http://www.bankofengland.co.uk/research/Pages/onebank/threecenturies.aspx.

### 3.5.2. The model

In the context of the New Keynesian model, the aggregate demand equation is derived from the intertemporal Euler equation and is forward-looking. The empirical literature has also used a backward-looking specification where the propagation mechanism is more appropriate when a new policy regime is introduced and the learning process about the new policy environment is gradually happening (Rudebusch & Svensson, 1999). During the sample period, the UK has established alternative policy regimes, such as fixed and floating exchange rates, and monetary and interest rate targeting regimes. In addition, the propagation mechanism matches pretty well the stylized facts referring to the persistent responses of real output to economic and policy disturbances (Fuhrer and Moore ,1995; Fuhrer and Rudebusch, 2004; Hafer and Jones, 2008; Rudebusch, 2002; Rudebusch and Svensson, 1999). In theory, a backward-looking specification is explained by a habit formation mechanism where the utility depends on current consumption relative to past consumption, and this relation adds significant persistent output dynamics (Fuhrer, 2000).

Initially, we estimate an aggregate demand model with constant Markov transition probabilities under the assumption that the mean growth rate of real output, the propagation mechanism, the lagged value of the debt-to-GDP ratio, and the error variance are subject to regime shifts:

$$y_t = a_s + \rho_s y_{t-1} + \beta_s d_{t-1} + \gamma' x_{t-j} + \sigma_s u_t \qquad (3.5.1)$$

$$p_{11,t} = Pr\left(s_t = 1 \middle| s_{t-1} = 1, \theta_1\right)$$

$$p_{22,t} = Pr\left(s_t = 2 \middle| s_{t-1} = 2, \theta_2\right)$$

where $y_t$ is the real output growth, $d_t$ is the debt-GDP ratio, $x_t$ is a vector containing the first difference of the real long-term interest rate, $\Delta r_t$, the inflation rate, $\pi_t$, and the log first difference of the real exchange rate, $e_t$. $p_{11}$ and $p_{22}$ are constant Markov transition probabilities, and $\theta$ is a vector with the estimated parameters of the model in each of the two regimes. The model allows for different sources of regime shifts, such as changes in the intercept, the dynamics of real output, the debt-to GDP ratio, and the error variance.[11] In this context, the two regimes are identified as high and low growth rates of real output. The lag order of the common regressors has empirically chosen using the SIC. The variable $d_{t-1}$ enters the equation with a negative sign, as we have discussed in the introduction. In the empirical literature (Juselius, Borio, & Disyatat, 2017), the aggregate demand depends on the difference between the actual and the equilibirum real interest rate, $r_{t-1} - r_{t-1}^{N}$. In the empirical analysis, the variable $\Delta r_{t-1}$ has replaced $r_{t-1} - r_{t-1}^{N}$ , having assumed that the equilibrium real interest rate depends on actual real interest rate of the last period (Couere, 2017). This variable enters the equation with a negative sign indicating the negative impact on consumption and consequently on real output growth. The variable $\pi_{t-1}$ enters the equation with a negative sign, as we have discussed in the introduction, and finally the variable $e_{t-1}$ enters the equation with a negative sign, showing that an appreciation of the pound sterling will depress aggregate demand and real output.

In turn, we estimate an aggregate demand model with time-varying Markov transition probabilities under the assumption that the mean growth rate of real output, the propagation mechanism and the error variance are subject to regime shifts:

$$y_t = a_s + \rho_s y_{t-1} + \gamma' x_{t-j} + \sigma_s u_t \qquad (3.5.2)$$

$$p_{11,t} = Pr\left(s_t = 1 \middle| s_{t-1} = 1, d_{t-1}, \theta_1\right)$$

$$p_{22,t} = Pr\left(s_t = 2 \middle| s_{t-1} = 2, d_{t-1}, \theta_2\right)$$

where $p_{11}$ and $p_{22}$ are time-varying transition probabilities which depend on the lagged value of the growth rate of the debt-to-GDP ratio.

[11] In a recent article, Arestis et al. (2016) estimate the impact of possible sources of regime shifts of output growth and inflation using different models.

The Bai-Perron test, used in the previous chapter, has been adopted here both to indicate the presence of nonlinearities, which the regime-switching specification could address. Thus, we estimate the first-order autoregressive aggregate demand model and have allowed the mean and the propagation mechanism to vary during the sample. The results reported in Table 3.2 show that there is a break in 1829. The mean and the propagation mechanism are statistically significant in the two sub-samples, and the different signs of the propagation mechanism indicate that the dynamics of the real economy have changed during the sample. It is woth noting that the break reflects the turning point in the industrial revolution of UK economy. Particularly, during the sub-sample 1760 to 1828 the UK economy has experienced the rise of the first industrial revolution where cotton textiles were the leading manufacturing products, while during the sub-sample 1829 to 2016, the economy has experienced the spread of industrialization which initially regarded transport innovations and the adoption of liberal trade policies through tariff reductions (World Development Report, 1987).

**Table 3.2: Bai-Perron estimates with breaks of real output growth**

| Samples | Constant | lagged output growth |
|---|---|---|
| 1762 - 1828 (67 obs.) | 1.751 ***<br>(0.362) | -0.358***<br>(0.097) |
| 1829- 2016 (188 obs.) | 1.475 ***<br>(0.304) | 0.291 ***<br>(0.088) |
| Adjusted $R$-squared = 0.10, SEE=2.84, Durbin-Watson=2.02 | | |
| *Sequential F-statistic determined breaks:* 1 | | |
| | $F$-statistic | *Scaled* $F$-statistic [CV] |
| 0 vs. 1 *** | 16.386 | 32.773 [11.47] |
| 1 vs. 2 | 4.862 | 9.724 [12.95] |
| *Break dates* | | |
| | Sequential | Repartition |
| 1 | 1829 | 1829 |
| **Notes**: Sample size: 257. *** indicates significance at 1%. Bai-Perron (2003) critical values (CV), obs=observations. | | |

Having identified the presence of nonlinearity in the behavior of the real output growth, we proceed to address it in the context of a regime-switching specification. In oder to explore the main source of the nonlinearity, we initially estimate two nested variants of the model (3.5.1) which assume fixed transition probabilities. The first one assumes that the mean and the dynamics change between the two states, and the second one assumes that the mean, the dynamics and the variance change between the two states.

The maximum lag length of the common regressors is set to be equal to four, and the value of SIC is equal to 5.2738. The Q-statistic at 36 lags is equal to 41.366 with a *p*-value=0.248 and shows the absence of serial correlation which is also verified for each time lag of the autocorrelation function at 5% significance level. Then, we have dropped the insignificant lags and re-estimated the autoregressive model. The value of SIC 5.106 is smaller than before and thus the selected model includes the following common regressors: the first and second lag of inflation, the first lag of the real interest rate, and the fourth lag of the growth rate of the real exchange rate. The Q-statistic of the model at 36 lags is equal to 41.446 with a *p*-value=0.245 and shows the absence of serial correlation which is also verified for each time lag of the autocorrelation function at 5% significance level.

Table 3.3 presents the results of the selected Markov switching model with common and error specific variance. The regressors that have been included in the regimes section are the constant, the lagged value of the growth rate of real GDP and the lagged value of the growth rate of debt-ratio. Two regimes are used to present the results which indicate the difference between the high and low growth periods.

In the upper part of the Table 3.3, we present the results with common variance. The following observations are made. First, the mean growth rate of real output in the high growth regime is 3.92%, while the mean growth rate of real output in the low growth regime is 0.72%. The Wald test for equal means across regimes does not reject the null hypothesis ($\chi^2(1)$=22.26, $p$= 0%) and thus the Markov switching model is appropriate to model changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. Second, the lagged value of the dependent variable is significant in both states and more persistent in the high growth state. This finding supports the empirical facts that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods and thus a nonlinear model is regarded more appropriate to capture these distinct patterns in the data. Third, the lagged growth rate of the debt-to GDP ratio has information content useful to predict movements in the growth rate of the real output in the high growth regime. The sign is negative indicating that as the public debt increases relative to GDP the growth rate of real output is expected to decrease. This result is consistent with the evidence documented in the empirical literature that increases in the public debt are detrimental to economic growth. Fourth, the common regressors are statistically significant and carry the correct signs. For instance, the change in the real interest rate has a negative effect on real output growth, suggesting that an increase in the real interest rate reduces aggregate demand and consequently the growth rate of real output. In addition, an appreciation of the real effective exchange rate reduces aggregate demand and consequently the growth rate of real output. The inflation rate initially reduces real output growth and then this negative effect is reversed, implying a short-lived negative effect. Fifth, the transition matrix parameter of $p_{11}$ is 0,798 implying a tendency of the economy to remain in the high growth state 1 when it is there, and the parameter of $p_{21}$ is -0.813 implying that the economy does not have a tendency to move from the low growth state 2 to high growth state 1. However, both parameters are not statistically significant. Sixth, the constant transition probabilities are 69% of remaining in each regime with the expected durations to be about three years in each regime.

In the lower part of the Table 3.3, we report the results with regime specific error variance. The following observations are made. First, there is a high and low growth regime with the mean growth rate of real output in the high growth regime 1 is 3.52%, while the mean growth rate of real output in the low growth regime 2 is 1.39%. The Wald test for equal means across regimes does not reject the null hypothesis ($\chi^2(1)$=16.609, $p$ =0%) and thus the Markov switching model is appropriate to model changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. Second, the volatility of the growth rate is higher in the low growth regime 2 due to a highly uncertain macroeconomic environment. The Wald test for equal volatilities across regimes does not reject the null hypothesis ($\chi^2(1)$=6.878, $p$=1%), thus indicating that the two regimes have different volatilities. Third, the lagged value of the dependent variable is significant in the high growth state and more persistent thus supporting the empirical facts that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods and consequently a nonlinear model is regarded more appropriate to capture these distinct patterns in the data. Fourth, the lagged growth rate of the debt-to GDP ratio has information content useful to predict movements in the growth rate of the real output in the high growth regime. The sign is negative indicating that as the public debt increases relative to GDP the growth rate of real output is expected to decrease. This finding is consistent with the evidence documented in the empirical literature about the negative growth effects of public debt. Fifth, the common regressors are statistically significant and carry the correct signs. Sixth, the transition matrix parameter of $p_{11}$ is 0.99 implying a tendency of the economy to remain in the low growth state 1 when it is there and it is statistically significant at 6% level, and the parameter of $p_{21}$ is -0.191 implying that the economy does not have a tendency to move from the high growth state 2 to low growth state 1 and it is not statistically significant. The constant transition probabilities are 73% of remaining in the low growth regime 1 and 55% of remaining in the high growth regime 2 with the expected durations to be about three and a half years in the first regime and two years in the second regime.

The implications of the empirical analysis of the Markov-switching model with fixed transition probabilities, and common and error specific variances are twofold. First, the presence of nonlinearities is related to different means, volatilities and dynamics of real output growth across samples. Second, the lagged value of the growth rate of the debt-to GDP ratio has predictive content useful to understand movements in the growth rate of the real output in the high growth regime. The negative sign shows that an increase in the public debt relative to GDP is expected to decrease the growth rate of real output.

**Table 3.3: Estimation results of Markov switching model with fixed transition probabilities**

| Error variances | Regimes | Constant | $y_{t\text{-}1}$ | $d_{t\text{-}1}$ | *Common regressors* $\Delta r_{t\text{-}1}$ | $e_{t\text{-}4}$ | $\pi_{t\text{-}1}$ | $\pi_{t\text{-}2}$ |
|---|---|---|---|---|---|---|---|---|
| *Common variance* | High growth **(regime 1)** | 3.92<br>(7.76)*** | -0.65<br>(-2.95)*** | -0.20<br>(-2.05)*** | -0.98<br>(-3.77)*** | -7.07<br>(-2.47)*** | -1.07<br>(-4.04)*** | 0.95<br>(3.64)*** |
| | Low growth **(regime 2)** | 0.73<br>(1.19) | 0.525<br>(3.78)*** | 0.025<br>(0.78) | | | | |
| LOG(SIGMA): 0.85 (12.25)***<br>Transition matrix parameters: $p_{11}$=0.80 (1.16), $p_{21}$= -0.81 (-1.03)<br>Transition probabilities: : $p_{11}$=0.69, $p_{22}$=0.69<br>*Wald Test* (*equal means*): 3.192 (0.677)***, $\chi^2$(1)-statistic = 22.26 [$p$-value=0%] | | | | | | | | |
| *Regime specific error variance* | High growth **(regime 2)** | 3.52<br>(10.35)*** | -0.31<br>(-2.51)*** | -0.06<br>(-3.83)*** | -0.88<br>(-6.10)*** | -9.27<br>(-5.17)*** | -0.90<br>(-7.00)*** | 0.87<br>(6.77)*** |
| | Low growth **(regime 1)** | 1.39<br>(2.78)*** | 0.12<br>(0.87) | -0.002<br>(-0.005) | | | | |
| LOG(SIGMA): **High growth (regime 2)**: -0.27 (-0.85), **Low growth (regime 1)**: 1.23 (13.62)***<br>Transition matrix parameters: $p_{11}$ = 0.99 (1.87), $p_{21}$ = -0.19 (-0.51)<br>Transition probabilities: : $p_{11}$=0.73, $p_{22}$=0.55<br>*Wald Test* (*equal means*): -2.13 (0.523)***, $\chi^2$(1)-statistic = 16.609 [$p$-value=0%]<br>*Wald Test* (*equal variances*) : 0.963 (367)***, $\chi^2$(1)-statistic = 6.878 [$p$-value=1%] | | | | | | | | |
| **Notes**: Sample size: 257. Numbers are estimated coefficients and numbers in parentheses are standard errors. Huber-White heteroskedasticity-consistent standard errors. *** indicate significance at 1% level. | | | | | | | | |

In the aggregate demand model (3.5.1), we have allowed for the lagged value of the growth rate of debt ratio to constitute one of the possible sources of the regime shift. Then, we have estimated the aggregate demand model (3.5.2) which allows the time-varying Markov transition probabilities to depend on the lagged value of the growth rate of debt ratio. In other words, the lagged movements in the leading indicator of the growth rate of real output influence the transition probabilities from period *t*-1 to period t. This type of modelling has been popularized by Filardo (1994). The non-switching regressors include lagged value of the change in the real long-run interest rate, the lagged values of the inflation rate, and the lagged value of the growth rate of the exchange rate.

The results are shown in the Table 3.4 with common and different error variances. In the upper part of this Table, we present the results with common variance. The value of the Q-statistic for the absence of serial correlation at 36 lags is equal to 38.821 with a *p*-value=0.344, and thus we cannot accept the presence of autocorrelation in the estimate model. The following observations can be made: First, the coefficients on the intercept in the mean equation have opposite signs, which correspond to the fast (regime 2) and slow (regime 1) growth rates for the U.K. economy, and they are statistically significant only in the second regime. The Wald test for equal means across regimes does not reject the null hypothesis ($\chi^2(1)$=12.04, *p*=0%), and thus the Markov switching model is appropriate to model changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. Second, the lagged value of the dependent variable is significant in both states and more persistent in the state of the slow growth state. This finding is not consistent with the empirical facts that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods. Thirdy, increases in the lagged growth rate of the debt ratio are associated with a higher probability of remaining in the contraction regime 1, and a higher probability of moving from the high growth regime 2 to low growth regime 1. Thus, the debt ratio constitutes a leading indicator of the transition probability across the two regimes. Fourth, the common regressors are statistically significant and carry the correct signs. Fifth, the transition probability matrix indicates that there is considerable state dependence in the transition probabilities with a relatively higher probability of remaining in the origin regime. In other words, we observe 63% probability for the low output regime, and 78% probability for the high output one. The corresponding expected duration in the first regime is approximately three years and in the second regime twenty nine years.

In the lower part of the Table 3.4, we report the results with regime specific error variance. The value of the Q-statistic for the absence of serial correlation at 36 lags is equal to 51.226 with a $p$-value=0.048, and thus we cannot accept the presence of autocorrelation in the estimated model. The following observations can be made: First, the coefficients on the intercept in the mean equation indicate a slow (regime 1) and fast (regime 2) growth rates for the U.K. economy and they are both statistically significant. The Wald test for equal means across regimes does not reject the null hypothesis ($\chi^2(1)$=40.35, $p$=0%), indicating that the two regimes have different means. Second, the lagged value of the dependent variable is significant only in the fast growth regime 2 and more persistent. This finding is consistent with the empirical facts that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods. Third, the volatility growth is statistically significant in both regimes, and is higher in the low growth regime due to a highly uncertain macroeconomic environment. The Wald test for equal volatilities across regimes does not reject the null hypothesis ($\chi^2(1)$=30.72, $p$=0%), indicating that the two regimes have different volatilities. Fourth, increases in the lagged growth rate of the debt ratio are associated with a higher probability of moving from the high growth regime 2 to low growth regime 1. Thus, the debt ratio constitutes a leading indicator of the transition probability across the two regimes. Fifth, the transition probability matrix indicates that there is considerable state dependence in the transition probabilities with a relatively higher probability of remaining in the origin regime only in the case of low regime. In other words, we observe 72% probability for the low output regime, and 47% probability for the high output one. The corresponding expected duration in the first regime is approximately four years and in the second regime five years.

The implications of the empirical analysis of the Markov-switching model with time-varyiung transition probabilities, and common and error specific variances are twofold. First, the presence of nonlinearities is related to different means, volatilities and dynamics of real output growth across samples. Second, the debt ratio constitutes a leading indicator of the transition probability across the two regimes. In particular, an increase in the growth rate of debt-to GDP ratio increases the probability of moving from the high growth regime to low growth regime, something which is consistent with the result obtained from the fixed transition probabilities where an increase in the public debt relative to GDP is expected to decrease the growth rate of real output. These findings fill in the gap in the literature on the link between real output growth and public debt, which constitutes an important topic from a policy perspective, and thus support the arguments for long-run fiscal consolidation.

**Table 3.4: Estimation results of Markov switching model with time-varying transition probabilities**

| | | *Shifting regressors* | | *Transition parameters* | | *Common regressors* | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | Constant | $y_{t-1}$ | $p_{11}$-$d_{t-1}$ | $p_{21}$-$d_{t-1}$ | $\Delta r_{t-1}$ | $e_{t-4}$ | $\pi_{t-1}$ | $\pi_{t-2}$ |
| *Common variance* | High growth **(regime 2)** | 3.508*** (0.477) | -0.282*** (0.106) | -0.109* (0.064) | 0.288* (0.155) | -0.98*** (0.254) | -6.506** (2.808) | -1.025*** (0.255) | 0.928*** (0.254) |
| | Low growth **(regime 1)** | -0.131 (0.726) | 0.595*** (0.166) | | | | | | |
| LOG(SIGMA): 0.862*** (0.071)<br>Transition matrix parameters: $p_{11}$-*constant* = 0.56 (0.499), $p_{21}$-*constant* = -1.799***(0.69)<br>Transition probabilities: : $p_{11}$=0.64, $p_{22}$=0.79<br>*Wald Test* (*equal means*): 3.377 (0.973)***, $\chi^2(1)$-statistic = 12.04 [*p*-value=0%] | | | | | | | | | |
| *Regime specific error variance* | High growth **(regime 2)** | 4.303*** ( 0.34) | -0.433*** (0.118) | 0.003 (0.055) | 0.332** (0.157) | -1.088*** (0.193) | -10.539*** (1.831) | -1.165*** (0.197) | 1.1038*** (0.183) |
| | Low growth **(regime 1)** | 1.025*** (0.435) | 0.202 (0.129) | | | | | | |
| LOG(SIGMA): **High growth (regime 2)**: 0.014 (0.142), **Low growth (regime 1)**: 1.183*** (0.08)<br>Transition matrix parameters*:* $p_{11}$-*constant*= 0.936 ***(0.345), $p_{21}$-*constant*= 0.643 (0.679)<br>Transition probabilities: : $p_{11}$=0.72, $p_{22}$=0.47<br>*Wald Test* (*equal means*): -3.278 (0.526)***, $\chi^2(1)$-statistic = 40.35 [*p*-value=0%]<br>*Wald Test* (*equal variances*) : 1.169 (0.21)***, $\chi^2(1)$-statistic = 30.72 [*p*-value=1%] | | | | | | | | | |
| **Notes**: Sample size: 257. Numbers are estimated coefficients and numbers in parentheses are standard errors. Huber-White heteroskedasticity-consistent standard errors. ******* indicate significance at 1% level. | | | | | | | | | |

## 3.6. Conclusion

Several empirical studies have examined the link between public debt and real output growth. The discussion of these studies reveals the gap in the literature refering to a non-linear analysis of the empirical fact that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods, and the role of the debt-to-GDP ratio in driving the transition probabilities between these dynamic structures, and this chapter asprires to fill this gap. In particular, this chapter adds to the literature by examining the information content of the growth rate of the debt-to-GDP ratio in understanding the asymmetric behavior over real output expansions and contractions, using annual data from the middle of eighteenth century until 2016. In addition, the chapter examines whether the growth rate of the debt-to-GDP ratio is a leading indicator of the time-varying transition probabilities between expansions and contractions of the real output growth. The importance of this topic stems from the fact that the link between debt and output growth is crucial from a policy perspective, as the establishment of negative growth effects of government debt will support the arguments for long-run fiscal consolidation.

We have specified a backward-looking aggregate demand equation which relates the growth rate of real output on its own past value, the past value of the debt-GDP ratio, and past values of the change in the real long-term interest rate, the inflation rate, and the log first difference of the real exchange rate. The model allows for different sources of regime shifts, such as changes in the intercept, the dynamics of real output, the debt-to GDP ratio, and the error variance. In this context, the two regimes are identified as high and low growth rates of real output.

Firstly, we have estimated the model with constant Markov transition probabilities under the assumption that the mean growth rate of real output, the propagation mechanism, the lagged value of the debt-to-GDP ratio, and the error variance are subject to regime shifts. The results with common variance indicate that the means are different across sample and thus the Markov switching model is appropriate to model changes in the historical growth rates of real output and understand the asymmetric behavior over real output expansions and contractions. The evidence that the lagged value of the dependent variable is significant in both states and more persistent in the high growth state supports the empirical facts that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods and thus a nonlinear model is regarded more appropriate to capture these distinct patterns in the data. The lagged value of the growth rate of the debt-to GDP ratio has predictive content useful to understand movements in the growth rate of the real output in the high growth regime. The negative sign shows that an increase in the public debt relative to GDP is expected to decrease the growth rate of real output. The results with error specific variance are similar with those obtained with common variance, and, in addition, we observe that the output volatily is different across the two regimes. The implications of the empirical analysis of the Markov-switching model with fixed transition probabilities, and common and error specific variances are twofold. First, the presence of nonlinearities is related to different means, volatilities and dynamics of real output growth across samples. Second, the lagged value of the growth rate of the debt-to GDP ratio has predictive content useful to understand movements in the growth rate of the real output in the high growth regime. The negative sign shows that an increase in the public debt relative to GDP is expected to decrease the growth rate of real output.

In turn, we have estimated the model with time-varying transmission probabilities, which depend on the lagged value of the growth rate debt-to-GDP ratio, and examined the hypothesis whether the lagged growth rate of the debt ratio is a leading indicator which influence the transition probabilities of real output growth from period $t$-1 to period $t$ between states of expansion and contraction. The results with common and error specific variances showed that increases in the lagged growth rate of the debt ratio are associated with a higher probability of moving from the high growth regime to low growth regime. Thus, the debt ratio constitutes a leading indicator of the transition probability across the two regimes.

In a nutshell, the conclusions which result from the Markov-switching analysis and answer the research question is that firsly the lagged growth rate of the debt ratio has predictive content useful to understand movements in the growth rate of the real output in the high growth regime, with the negative sign showing that an increase in the ratio is expected to decrease the growth rate of real output, and secondly an increase in the lagged growth rate of the debt ratio is associated with a higher probability of moving from the high growth regime to low growth regime, and thus the ratio constitutes a leading indicator of the transition probability across the two regimes. These results fill in the gap in the literature on the link between real output growth and public debt and support the arguments for long-run fiscal consolidation.

# Chapter 4

# Modelling the Trade Flows between UK and Eurozone: A Global VAR analysis

## 4.1. Introduction

On the 31st of January, 2020, the UK withdrew from the European Union (EU) after almost half a century of membership. The process of Brexit (leaving the EU) had begun with the outcome of Britain's referendum, which took place on the 23rd of June, 2016. The decision to leave the EU has spurred a discussion about the potential risks for the UK economy, since the EU constitutes its largest trading partner. In 2016, for instance, when the referendum took place, UK exports of goods and services to the other EU member states were £236 billion and accounted for 43% of all UK exports, while the UK imports of goods and services from the other EU members states were £312 billion which amounted to about 54% of all imports, and thus the UK registered an overall trade deficit of £82 billion. In addition, the trade in services amounted to 38% of the overall exports to the EU (Ward, 2017). Germany was the largest origin of imports amounted to £75.12 billion and the second largest destination for exports amounted to £49.12 billion, whereas the USA is the largest destination for exports amounted to £99.57 billion and the second largest origin of imports amounted to £66.31 billion (Office for National Statistics, 2018). According to the statistics published by the Department for International Trade in 2018, among the top UK export and import markets were four big EU countries – Germany, France, Italy and Spain – and the USA. In particular, 21% of the UK exports of goods and services were directed to the four big EU countries and 18,8% to the

USA, while 26,5% of the UK imports for goods and services were come from the four big EU countries and 11,4% from the USA.

Given the fact that the EU constitutes the largest trading partner of the UK, Brexit would affect the UK economy through the trade channel. Thus, an interesting research question is whether the historical trade relations between the UK and the EU provide information about the potential effects of Brexit on the UK and the EU economies. In the empirical literature, there exist a few papers which examine the potential effects of the Brexit on the UK economy, using structural and gravity models. The results derived from these studies seem to suggest that the potential costs could be substantial. In particular, Baker et al. (2016), using the National Institute Global Econometric Model, have forecasted that the exit of the UK from the EU would be harmful for its GDP while at the same time a high inflation would be observed due to the effective pound exchange rate being depreciated. Dhingra et al. (2016), using a trade model of the global economy, have estimated that income will be decreased by between 1.3% and 2.6% due to trade being affected by Brexit and this fall will rise to between 6.3% and 9.5% as long as the long run effects of Brexit on productivity are evaluated. Gudgin et al. (2017), using a gravity model, have estimated that the overall impact of Brexit would be relatively small compared to the estimations of the Treasury, after accounting of the potential effects of the deprecation of sterling on exports which are expected to last for the next decade. Arregui and Chen (2018), using a computable general equilibrium model, have examined whether trade barriers between the UK and the EU would affect the real output in the UK. They have found that if the trade is free, then the real output would fall by an average of 3%, whereas if the trade is subject to barriers, then the real output would fall by an average of 6% .

Although most studies suggest that the potential costs of Brexit to the UK economy could be substantial, the literature is rather limited, and this requires more research to examine the significance and duration of the potential risks for the UK and the EU economies. This chapter aspires to fill in this gap in the literature by using an alternative approach to model the trade interactions between the UK, the four big euro area countries – Germany, France, Italy and Spain – and the USA, and forecast whether these historical trade flows provide information about the impact of Brexit on all countries in the analysis. The alternative approach we use to answer the research question is the innovative tool of the global vector autoregression (GVAR, Pesaran et al., 2004). In the context of this approach, we jointly model the behavior of global real imports and exports, as these variables co-vary together over the business cycle, because of the strong import component of exports due to globalization and internationalization of production (Bussiere et al., 2009). Economic theory postulates that real imports and real exports depend on real income and the real exchange rate. Thus, we set up a four-equation trade flows VAR specification, consisting of real imports, real exports, real income and real effective exchange rate. This specification is estimated as a GVAR which includes an explicit model for each of the six countries. Then, we use the model to perform impulse response analysis and forecast error variance decomposition. In particular, we forecast the effects of a negative shock to the UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area, on the economy of all countries in the analysis. In addition, we compute the proportion of the forecast variance of real imports and real exports of the UK and Germany that is explained by shocks to all variables and countries in the model. This research is important to assess the potential effects of Brexit on the economy of all countries in the analysis and the duration of these effects.

The rest of the chapter is organized as follows. Section 4.2 overviews the empirical literature. Section 4.3 presents the GVAR method. Section 4.4 presents the empirical analysis. Section 4.5 concludes.

## 4.2. Literature Review

The studies we have reviewed in the previous section suggest that the potential costs of Brexit to the UK economy could be substantial. However, the literature is rather limited and this requires more research to examine the significance and duration of potential risks on UK and EU economies. This chapter aspires to fill in this gap in the literature. In this section, we review some papers which use the GVAR approach to examine global economic interactions.

Dees, Di Mauro, Pesaran and Smith (2007) studied vector error correction models in an individual country framework and including domestic variables which are linked to country-specific foreign ones. They used data from 26 countries, while the euro area was treated as a single country with the dataset being between 1979 and 2003. The USA is also found in the model in order to incorporate the global oil price. In terms of variables, they included short- and long-run interest rates, real GDP, inflation rate, real equity prices and real exchange rate. The results showed that shocks from the financial sector were transmitted quite quickly from the Unites States into the euro area. The oil price shock affected significantly the inflation, but not the output. Finally, regarding the shocks from USA monetary policy, the effects on euro area were not significant. Hiebert and Vansteenkiste (2007) worked on how trade openness and technological shocks affected the manufacturing industrial market of the USA. A rise in the trade participation of the developing countries and fast technological advance were the main motivations behind this research. In this paper, there were no individual countries examines, but instead they focused solely on the industry of the United States. More specifically, they extracted data from 12 different sectors which were related to manufacturing process and the data period was between 1977 and 2003. The variables used were productivity, stock of capital, full-time equivalent employment and real compensation per employee. Two common exogenous variables were used as global ones, which were the oil price and the expenditure for research and development per employee. The choice of the GVAR method was due to the examination of the effects of exogenous variables' shocks as well as the potential spillovers originated from shocks on employment. The primary results showed that shocks on technology had greater effects than shocks on trade openness. Furthermore, the effect of trade openness was mainly negative but not significant on employment and compensation, while the technological progress had a positive and important impact on them. Regarding the spillover effects, these were found to be positive for the productivity and employment. Bussiere, Chudik and Sestieri (2009) studied the imbalances and the flows of exports and imports in a global level. Their contribution is twofold: firstly, they applied a GVAR model as it is the most suitable one for the examination of global linkages and secondly, they modeled the exports and imports as a joint variable instead of two separate ones as been done by the majority of trade papers due to the existence of the international control of production links. For their research, they used data between 1980 and 2007 for 14 developed countries and 7 emerging ones. In this paper, the euro area is not taken as a joint case and instead France, Germany, the Netherlands, Spain and Italy are included. The main variables that were used are exports, imports, real exchange rate and GDP plus the global variable of oil price. They also added foreign series specific to each of the four countries which correspond to cross section averages of the above main variables. With their empirical model, they observed that firstly, when there was a shock to the American output, this had an effect on exports mainly from non-European countries. Secondly, a negative shock to the real exchange rate of the US had a positive effect on exports from other countries but the degree of the latter effect was smaller than the one regarding the output. Changes to German and Chinese variables also affected significantly the other countries. Furthermore, they made a forward step to check whether they could predict the big financial crisis of 2008 by using the growth rates of exports and imports across their 21 countries. They found out that the fall in imports could be successfully predicted while for the imports there was an underestimation. Galesi and Sgherri (2009) studied the spillover effects of a negative shock of the American financial system into Europe. For this study, 27 countries in total were used which were 17 developed and 9 developing countries from Europe as well as the United States and the dataset was dated between June 1999 and April 2008. As this research focused on the financial sector, the variables that were contained were the real credit to enterprises, the growth rates of real equity prices, the real interbank rate and the real GDP. Regarding the foreign variables, these were constructed by taking into account the financial weights which reflected the bank lending relations between the examined countries. The importance of the use of GVAR methodology was on the fact that a group of financial and macroeconomic variables as well as international banking flows could be taken into consideration at the same time. The principal empirical results showed that the transmission of financial crisis from the USA into Europe was pretty quick thanks to asset prices mainly in the short-run, while in the long-run the channel which was responsible for this event was the credit quantity and cost. Feldkircher and Korhonen (2012) examined the case of China and its interactions with the global economy. During the past few decades, China emerged as a strong economy thanks to its huge presence in investment and world trade. It did not also suffer from the recent global crisis of 2008, unlikely with most of the developed countries around the world. To investigate the effects of changes of the Chinese output, the American output and the global oil on the Chinese economy, the GVAR methodology was used as it was the most suitable one to take into account a large number of countries at the same time and shocks coming from macroeconomic variables. In this study, a total of 52 economies were taken into account where the euro area was considered as a single economy. This group of economies was a mix of developed, large and small developing countries coming from almost all the continents. In addition, the dataset was dated between 1995 and 2011 and the variables included were the inflation, the real GDP, the real exchange rate as well as the short- and long-term interest rates. However these rates could not be obtained for all the examined countries, mostly for the cases of the emerging ones. Finally, the global variable of oil price was also included. With respect to the empirical results, they found out that the country which benefitted from a rise in the Chinese output was Brazil, while the Asian countries had the smallest benefits. In order to compare China with a larger economy such as the United States, a positive shock of the American output was triggered to examine its effects and they observed that its larger trading allies such as Mexico and the UK got the highest increases in their outputs, while China was not affected almost at all. Finally, a rise in the oil price of China had a significant and positive effect mostly on Russia as it exports large amounts of oil globally while the effect on the Chinese output was negative. Cakir and Kabundi (2013) investigated the links of trade between South Africa and Brazil, Russia, India and China, which constitute the so-called BRIC group. The choice of South Africa was mainly due to the high volume of its export growth, especially when it started to trade more with the BRIC group rather than the USA and the EU. Furthermore, in 3 decades time, the growth of the BRIC countries will overpass the economic growth of the present developed ones. The GVAR was used here as it proved to be the only methodology that can take into account the international interactions and the country-specific shocks taking place in the whole world. They used data from 32 countries from all the continents for the period between 1995 and 2004. About the empirical part, two different estimations were applied for their research, as they split the countries into different groups each time. For the first one, the 8 euro area countries were considered as one join country with the rest 24 countries being considered separately and for the second one, the difference was the creation of an another joint country which consisted the BRIC countries. As a result for the latter estimation, there were 2 joint countries and 20 individual ones. The variables which were obtained for this analysis were the real exports, the real imports, the real output, the real effective exchange rates and the inflation, as all of these affect trade in a direct way. The oil price was also included as a common factor for all the countries. The main empirical results showed that taking into account the generalized impulse response functions, the shocks that came from the exports of the BRIC group affected significantly the imports and the output of South Africa, which proved the strong links between the BRIC countries and South Africa mostly on exports rather than on imports. Sun, Heinz and Ho (2013) studied the links among 33 countries across Europe by implementing both financial and trade weights inside their model in order to take into account the financial and trade relationships of the Central, Eastern and South-Eastern countries (CESEE) with the Western ones. The dataset was between 2000 and 2011 and the variables used were inflation, real GDP, real credit growth and long-term interest rate. They found out that there were significant effects on domestic GDP growth and long-run interest rates. On the other hand, the effects were weak regarding the inflation and real credit growth. Furthermore, the euro area countries seemed to be the group which affected the rest of Europe significantly, whereas the effects originated from the CESEE and Nordic countries were relatively smaller. Konstantakis and Michaelides (2014) investigated how the debt crisis of the past decade affected the United States and the group of EU15. More specifically they were interested in the direction in which this crisis was transmitted between the aforementioned countries. The GVAR approach was used for this purpose as it includes a modeling structure which can be applied globally and to incorporate the shocks applied to various variables and the channels through which these are transferred and affect the examined countries. For this study, the number of countries used was two, as the EU15 group was considered as one entity whereas the other one was the United States. The data period was between 2000 and 2011 and two variables were used for this analysis, the GDP and the debt ratio. Regarding the empirical results, they observed firstly the vulnerability of the EU15 group on shocks coming from the United States, however these effects seem to happen in the short-term only. About the debt crisis, the transmission direction is found to happen from the United States towards the European countries, while the opposite one does not exist. Finally, with respect to the GDP, the American GDP brings positive effect into the European one, while the effect of the European GDP towards the American one is found to be negative. Ricci-Risquete and Ramajo-Hernandez (2014) studied how shocks of macroeconomic variables which happen in a country member of the European Union affect the rest of the union. This topic is important from the economic policy view as European countries need to know which policy to implement after observing changes to fiscal elements in neighboring ones. For this purpose, 14 countries from the EU15 group with the exception of Luxembourg as well as the United States were included in this model with the dataset dating between 1978 and 2009. The choice of these countries was done according to data availability, their contribution to the global GDP and statistical purposes. Regarding the variables, they used the GDP, the total public receipts, the consumer price index, the total public expenditure, the exchange rate as well as both the short- and long-run interest rates. The oil price is included as a global variable as well. After applying shocks to fiscal variables mostly in France and Germany, they found out that the effects are greater domestically rather than being transmitted into other countries. Furthermore, when totals government receipts rise and total government expenditure declines, these contribute positively to the GDP. In additional, when shocks are applied globally rather than domestically, the effects seemed to be the same as the domestic ones regarding the GDP. In terms of policy coordination and execution, they suggested this process to take place in a slow pace in order the European countries to avoid experiencing negative economic results. Chisiridis, Mouratidis and Panagiotidis (2018) addressed the issue of trade asymmetry between the north euro and south euro areas as one of the main factors which triggered the 2008 financial crisis. They examined the effects coming from shocks to domestic macroeconomic variables such as non-export real output, real imports, real exchange rate as well as to oil supply which acts as a global variable. They extracted data for 28 countries which are dated between 1980 and 2016. The main reasons of incorporating two main European groups which are the north and the south are that in the south, government debt, inflation and unemployment are significantly higher than in the north and also the discrepancies in their current accounts. The empirical analysis showed that firstly positive shocks to demand affect positively the exports and investment, however current accounts are affected negatively or stay the same. On the other hand, when the real effective exchange rate was devaluated in the south area, the exports increased while there was no effect on imports. Faryna and Simola (2018) worked on the effects of shocks to foreign output and oil price on the Commonwealth of Independent States which was derived by the dismissal of the Soviet Union. The purpose of this study was to investigate the response of this group of countries to the interaction with the global economy. Moving from central planning economies to more open ones was not easy for those countries, as economic deceleration and high levels of inflation were observed. Their model included five ex-Soviet countries which comprised the Commonwealth group, with Russia being treated independently due to its size and 23 other countries across the world, including the euro area which is taken as a single country. The dataset was dated between 2001 and 2016 and the variables used were the real output, the consumer price inflation, the real exchange rate and the nominal short-run interest rate. Oil price was also included as a global variable. The main empirical results proved that the ex-Soviet countries seemed to be vulnerable to all types of shocks, both domestic and global ones. The biggest effects came from shocks from the area itself, as well as Russia and the USA. As the ex-Soviet group incorporates countries which do not affect the global economy significantly, their policymakers should keep an eye on economic progress happening regionally and globally in order not to be sensitive to shocks.

## 4.3. Empirical Methodology

In this chapter, we use the GVAR approach to forecast the effects of a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area, on the economy of all countries in the analysis. In addition, we compute the proportion of the forecast variance of real imports and real exposts of the UK and Germany that is explained by shocks to all variables and countries in the model.

An alternative approach to the GVAR is the Factor-Augmented VAR model which is useful if we estimate the dynamic responses of a lager number of domestic variables to disturbances in foreign countries (Mumtaz and Surico, 2009), which is not, however, the case in the present analysis.

We start the analysis of the GVAR approach by modeling each of the six economies as a country-specific one-lag VAR specification (VARX*), including four domestic variables – real imports *z*, real exports, *x*, real output *y*, and real effective exchange rate *e* – and four foreign variables,

$$x_{j,t} = \Gamma_0 + \Gamma_1 x_{j,t-1} + \Theta_{j,0} x^*_{j,t} + \Theta_{j,1} x^*_{j,t-1} + \zeta_{j,t} \tag{4.3.1}$$

where $j$=1,.., 6, $x_{j,t} = \left(z_{j,t}, x_{j,t}, y_{j,t}, e_{j,t}\right)'$ is a 4x1 vector of endogenous domestic variables of country $j$, $x^*_{j,t} = \left(z^*_{j,t}, x^*_{j,t}, y^*_{j,t}, e^*_{j,t}\right)'$ is a 4x1 vector of foreign variables, $\Gamma_0$ is a 4x1 vector of constant terms, $\Gamma_1$ is a 4x4 matrix of coefficients of lagged domestic variables, $\Theta_0$ is a 4x4 vector of coefficients of contemporaneous foreign variables, $\Theta_1$ is a 4x4 matrix of coefficients of lagged foreign variables, and $\zeta_t$ is a 4x1 vector of reduced-form real shocks. This model can be written as a vector error correction specification (VECMX*),

$$\Delta x_{j,t} = \alpha_j ECT_{j,t-1} + \Lambda_{j,0} \Delta x^*_{j,t} + \Lambda_{j,1} \Delta x^*_{j,t-1} + v_{j,t} \tag{4.3.2}$$

where $\alpha_j = \left(\alpha_{j,z}, \alpha_{j,x}, \alpha_{j,y}, \alpha_{j,e}\right)'$, $\Lambda_0$ is a 4x4 vector of coefficients of contemporaneous changes in foreign variables, $\Lambda_1$ is a 4x4 matrix of coefficients of lagged changes in foreign variables. The term ECT is the long-run co-integrating relation between the four domestic and four foreign variables (assuming the existence of one cointegrating vector), given by $ECT_{j,t-1} = \beta' * x_{j,t-1} + \beta^{*'} * x^*_{j,t-1}$. The rank of $\alpha_j \beta'_j$ is determined using the trace and maximum eigenvalue statistics, where $\beta_j = \left(\beta, \beta^*\right)'$ with $\beta = \left(\beta_{j,z} \quad \beta_{j,x} \quad \beta_{j,y} \quad \beta_{j,e}\right)'$ and $\beta^* = \left(\beta^*_{j,z} \quad \beta^*_{j,x} \quad \beta^*_{j,y} \quad \beta^*_{j,e}\right)'$. We can test the hypothesis of weak exogeneity of the foreign variables with respect to the domestic variables in each country-specific model. In the model of country $j$, we run the following auxiliary equation for foreign imports,

$$\Delta z^*_{j,t} = \phi_0 + \phi_1 ECT_{j,t-1} + \sum_{k=1}^{p} \theta_{j,z} \Delta z_{j,t-k} + \sum_{k=1}^{p} \psi_{j,z} \Delta \tilde{z}^*_{j,t-k} + \zeta_{j,t} \tag{4.3.3}$$

where, $\Delta\tilde{z}_j^* = \left( \Delta z_j^*, \Delta x_j^*, \Delta y_j^*, \Delta e_j^* \right)$, and test the hypothesis that $\phi_1 = 0$ , using the $F$-statistic. If the null hypothesis cannot be rejected, then the weak exogeneity is valid. The same procedure is followed for all foreign variables in each country-specific model.

After estimating each VARX* model for all countries of the sample, we proceed to specify and estimate the GVAR model. Let us, firstly, define the weigh matrix $W_j$ reflecting the trade structure between the country $j$ and the other five countries,

$$W_j = \begin{bmatrix} 0 & 0 & 0 & 0 & w_{0z1} & 0 & 0 & 0 & w_{0z2} & 0 & 0 & 0 & w_{0z3} & 0 & 0 & 0 & w_{0z4} & 0 & 0 & 0 & w_{0z5} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & w_{0x1} & 0 & 0 & 0 & w_{ox2} & 0 & 0 & 0 & w_{ox3} & 0 & 0 & 0 & w_{ox4} & 0 & 0 & 0 & w_{ox5} & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & w_{oy1} & 0 & 0 & 0 & w_{oy2} & 0 & 0 & 0 & w_{oy3} & 0 & 0 & 0 & w_{oy4} & 0 & 0 & 0 & w_{oy5} & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & w_{oe1} & 0 & 0 & 0 & w_{oe2} & 0 & 0 & 0 & w_{oe3} & 0 & 0 & 0 & w_{oe4} & 0 & 0 & 0 & w_{oe5} \end{bmatrix}$$

secondly, define the selection matrix $S_j$ ,

$$S_j = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

and finally define a 24x1 vector of all endogenous variables $x_t = x_{j,t}$ . Then, $x_{j,t} = S_j \times x_t$ , and $x_{j,t}^* = W_j \times x_t$ . Using the matrices $W_j$ and $S_j$ , the model (4.3.1) is written as

$$S_j x_t = \Gamma_0 S_j + \Gamma_1 S_j x_{t-1} + \Theta_{j0} W_j x_t + \Theta_{j1} W_j x_{t-1} + \zeta_{j,t} \tag{4.3.4}$$

and re-arranging yields,

$$\left( S_j - \Theta_{j0} W_j \right) x_t = \Gamma_0 S_j + \left( \Gamma_1 S_j + \Theta_{j1} W_j \right) x_{t-1} + \zeta_{j,t} \tag{4.3.5}$$

or,

$$\Psi_j x_t = \Xi_j + \mathrm{M}_j x_{t-1} + \zeta_{j,t} \tag{4.3.6}$$

where, $\Psi_j = S_j - \Theta_{j0} W_j$ , $\Xi_j = \Gamma_0 S_j$, $\mathrm{M}_j = \Gamma_1 S_j + \Theta_{j1} W_j$ .

Stacking all country-specific models, we get the GVAR model,

$$\Psi x_t = \Xi + \mathrm{M} x_{t-1} + \zeta_t \tag{4.3.7}$$

or, $$x_t = \Pi_0 + \Pi_1 x_{t-1} + \mu_t \tag{4.3.8}$$

where, $\Pi_0 = \Psi^{-1}\Xi$, $\Pi_1 = \Psi^{-1}\mathrm{M}$, $\mu_t = \Psi^{-1}\zeta_t$, $\sum_\mu = E[\mu_t \mu_t']$.

The moving average representation of the GVAR model is,

$$x_t = d_t + \sum_{s=1}^{\infty} \mu_{t-s} \tag{4.3.9}$$

where $d_t$ is the deterministic component of $x_t$ and $A_s = \Pi_1 A_{s-1}$ .

In the context of the GVAR, we conduct the generalized impulse response function (GIRF) analysis which examines the effects of variable-specific disturbances on all the variables of the system. This method does not identify the shocks according to economic theory or a triangular form, but instead assumes that historical correlations of shocks are given (Chudik and Pesaran, 2014). The GIRF of one standard deviation innovation at time $t$ to the $k^{th}$ equation of the $l^{th}$ variable at time $t+h$ is given by

$$GIRF(x_t, \zeta_{kt}, h) = \frac{\upsilon_l' A_h \Psi^{-1} \sum_\zeta \upsilon_k}{\sqrt{\upsilon_l' \sum_\zeta \upsilon_k}}$$

where $k, l = 1,..4$, $h = 0,1,\ldots$, $\Omega$ is the information set available at time $t$, $\sum_\zeta$ is the variance-covariance matrix, $\upsilon_k$ is a selection vector with one as the $k^{th}$ element in a country-specific disturbance. We have used the GIRF approach to simulate a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area.

We also present the generalized forecast error variance decomposition (GFEVD) of disturbances to specific variables, given by

$$GFEVD(x_{lt}, \zeta_{kt}, h) = \frac{\sigma_{kk}^{-1} \sum_{h=0}^{n} \left( \upsilon_l' A_h \Psi^{-1} \Sigma_\zeta \upsilon_k \right)^2}{\sum_{h=0}^{n} \upsilon_l' A_h \Psi^{-1} \Sigma_\zeta \Psi^{-1'} A_h' \upsilon_k}$$

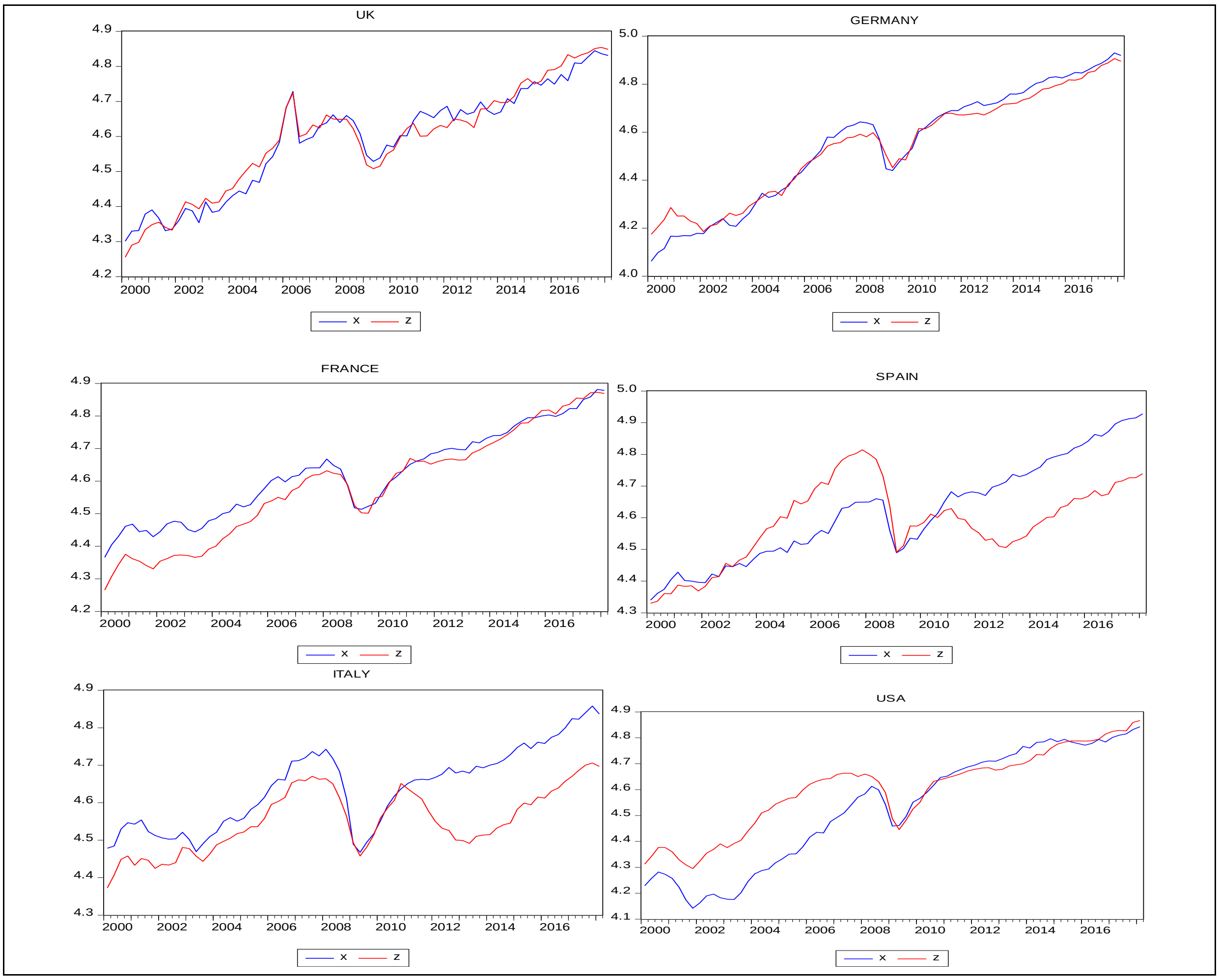
where $\sigma_{kk}$ is the diagonal element of the variance-covariance matrix $\Sigma_\zeta$ related to the $k^{th}$ equation in the $k^{th}$ country. This expression provides the fraction of the *h*-step ahead forecast error variance of the $l^{th}$ element of $x_t$ that is explained by conditioning on current and future values of the generalized innovations in the $k^{th}$ element of $x_t$.

# 4.4 Empirical Analysis

## 4.4.1 Data

We have used a sample period from the first quarter of 2000 to the first quarter of 2018, to reflect the initiation of the Euro during the last 2 decades, for the following variables: real imports, real exports, real GDP, and the real effective exchange rate. All the data were transformed to logarithms ans were obtained from Federal Reserve Economic Data St. Louis FRED. Figure 4.1 plots real imports and exports for all countries of the sample. On inspection we observe that these two variables co-vary together over the business cycle, thus justifying their joint modeling in the context of the GVAR.

Figure 4.1: Real imports and real exports of the six countries of the sample




## 4.4.2. Trade weights

Table 4.1 presents the weight matrix which has been used to construct the foreign variables and was based on the bilateral trade relations between the six countries of the sample. The bilateral trade between the euro area countries and the other countries has been obtained from EUROSTAT database of international trade of goods,[12] and the bilateral trade between the

12 EUROSTAT, http://ec.europa.eu/eurostat/web/international-trade-in-goods/data/database

USA and the other countries has been obtaines from US Census Bureau.[13] The trade weights have been constructed as the share of imports and exports of each country with its trading partners to its total trade during the period 2012-2017. The UK trade with the four euro area countries amounts to about 23%, and the higher share is with Germany, amounting to 11%, followed by the USA, amounting to about 10% which is close to 17% provided by Bussière et al (2009). France has the highest share of Germany's trade, amounting to 6,4%, followed closely by USA with a share of 6,2%. Germany has the highest share of French's trade, amounting to 16%, followed by Italy with a share of about 7%. France and Germany have the highest shares of Spain's trade, amounting to 12% and 11% respectively, and these two countries have the highest shares of Italy's trade, amounting to 9% and 13% respectively. The USA trade with the four euro area countries amounts to about 9%, and with the UK amounts to 3% which is close to 4,6% provided by Bussière et al (2009). For the inter-euro are trade, our figures are smaller than those provided by Bussière et al (2009. The figures we have calculated reflect the real-world counterparts, but they are small and do not sum up to one because the bilateral trade of each country of the sample is a small share compared to its total trade. However, we have adjusted these figures to sum up to one. The information which is derived from the trade matrix is that there exist trade interrelations between the countries of the sample which play key role in the transmission of real shocks between these countries.

**Table 4.1: Weight matrix based on trade relations (%)**

| | UK | GE | FR | SP | IT | US |
|---|---|---|---|---|---|---|
| UK | 0 | 0.046 | 0.047 | 0.049 | 0.039 | 0.029 |
| GE | 0.1143 | 0 | 0.157 | 0.109 | 0.133 | 0.044 |
| FR | 0.056 | 0.064 | 0 | 0.122 | 0.09 | 0.02 |
| SP | 0.028 | 0.025 | 0.062 | 0 | 0.046 | 0.006 |
| IT | 0.03 | 0.042 | 0.065 | 0.062 | 0 | 0.016 |
| US | 0.096 | 0.062 | 0.053 | 0.048 | 0.055 | 0 |

[13] https://www.census.gov/foreign-trade/balance/c4120.html

### 4.4.3 Unit roots

We examine the stationary properties of the data, by using the Augmented Dickey-Fuller (ADF) test proposed by Dickey and Fuller (1979), and the weighted symmetric version of the ADF test proposed by Park and Fuller (1995). Table 4.2 reports the unit root results for both domestic and foreign variables. First, considering the log-levels of all variables, we can see that both tests cannot reject the null hypothesis of a unit root, implying that the time-series concerned are integrated of order one (I(1)) processes. Then, considering the first difference of these variables, we conclude that the growth rates of the time-series concerned are stationary processes.

**Table 4.2: Unit root tests**

| | | | Domestic variables | | | | | | | Foreign variables | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | ST | 5% | UK | GE | FR | SP | IT | US | | UK | GE | FR | SP | IT | US |
| ***z*** | ADF | -2,89 | -1.26 | -0.54 | -0.56 | -1.96 | -2.01 | -0.98 | ***z**** | -0.89 | -1.08 | -1.02 | -0.80 | -0.85 | -0.89 |
| ***z*** | WS | -2,55 | 0.67 | 0.68 | 0.85 | -0.87 | -1.34 | -0.18 | ***z**** | 0.10 | 0.08 | 0.17 | 0.31 | 0.35 | 0.52 |
| **Δ*z*** | ADF | -2,89 | -6.85 | -4.10 | -5.01 | -4.54 | -4.40 | -4.83 | **Δ*z**** | -4.55 | -4.61 | -4.33 | -4.38 | -4.48 | -4.55 |
| **Δ*z*** | WS | -2,55 | -7.02 | -4.35 | -4.56 | -4.73 | -4.16 | -4.82 | **Δ*z**** | -4.46 | -4.51 | -4.32 | -4.63 | -4.42 | -4.50 |
| ***x*** | ADF | -2,89 | -1.07 | -1.11 | -0.46 | -0.33 | -1.62 | -0.49 | ***x**** | -0.77 | -0.65 | -0.89 | -0.86 | -0.74 | -0.91 |
| ***x*** | WS | -2,55 | 0.26 | 0.60 | 0.58 | 0.74 | -1.30 | -0.16 | ***x**** | 0.13 | 0.10 | 0.23 | 0.21 | 0.45 | 0.47 |
| **Δ*x*** | ADF | -2,89 | -6.39 | -4.71 | -4.43 | -5.56 | -4.31 | -5.05 | **Δ *x**** | -4.88 | -4.77 | -4.62 | -4.37 | -4.60 | -4.38 |
| **Δ*x*** | WS | -2,55 | -6.59 | -4.85 | -4.19 | -5.73 | -4.38 | -5.14 | **Δ *x**** | -4.95 | -4.81 | -4.74 | -4.41 | -4.67 | -4.48 |
| ***y*** | ADF | -2,89 | -1.09 | -0.08 | -0.82 | -1.49 | -2.39 | -0.01 | ***y**** | -0.32 | -1.26 | -0.63 | -0.62 | -0.42 | -0.74 |
| ***y*** | WS | -2,55 | 0.53 | 0.31 | 1.15 | -0.30 | -2.29 | 1.68 | ***y**** | 0.81 | 1.17 | 0.56 | 0.55 | 0.84 | 0.53 |
| **Δ*y*** | ADF | -2,89 | -4.13 | -4.49 | -3.67 | -1.96 | -4.36 | -4.03 | **Δ *y**** | -4.03 | -3.51 | -3.87 | -3.89 | -3.86 | -3.76 |
| **Δ*y*** | WS | -2,55 | -4.37 | -4.70 | -3.66 | -1.92 | -4.10 | -4.08 | **Δ *y**** | -4.14 | -3.56 | -4.02 | -4.03 | -3.99 | -3.93 |
| ***e*** | ADF | -2,89 | -1.59 | -1.45 | -1.32 | -2.76 | -2.02 | -1.76 | ***e**** | -0.97 | -0.86 | -1.02 | -0.96 | -0.88 | -1.22 |
| ***e*** | WS | -2,55 | -1.23 | -1.79 | -1.68 | -0.51 | -1.80 | -1.85 | ***e**** | -1.43 | -1.15 | -1.44 | -1.31 | -1.29 | -1.43 |
| **Δ*e*** | ADF | -2,89 | -5.30 | -4.39 | -5.14 | -5.04 | -4.31 | -5.61 | **Δ *e**** | -4.06 | -4.42 | -4.83 | -5.00 | -4.69 | -5.65 |
| **Δ*e*** | WS | -2,55 | -5.46 | -4.67 | -5.08 | -5.25 | -4.59 | -5.68 | **Δ *e**** | -4.21 | -4.58 | -4.80 | -4.88 | -4.65 | -5.58 |

**Notes:** Sample size: 73. ADF regressions with a constant and without trend. WS= weighted symmetric estimation of ADF equations. ** indicates significance at 5%

### 4.4.4. The VARX* model and cointegrating relations

For each country, we have estimated a vector autoregressive model with domestic and foreign variables with lag length of $p_j$ and $q_j$ respectively, assuming that the foreign variables are weakly exogenous. The domestic variables included in each country-specific model are the log-levels of the real imports, $z$, real exports, $x$, real income, $y$, and the real effective exchange rate, $e$, while the country-specific corresponding foreign variables, such as the foreign real imports, $z^*$, the foreign real exports, $x^*$, the foreign real income, $y^*$, and the foreign real effective exchange rate, $e^*$. The maximum lag order of both domestic and foreign variables is equal to 4, and the SIC has indicated that the optimal order of domestic and foreign variables is one, and thus a VARX*(1,1) for each country is selected. This finding is presented in Table 4.3, together with the number of cointegrating vectors derived from the trace statistics reported in Table 4.4 with the critical values at 5% significance level. For UK, Italy and USA, the analysis has indicated the presence of two cointegrating relations between the four variables. In Germany and France we observe one cointegrating relations, and in Spain we have obtained three cointegrating vectors.

We have estimated long-run trade equations of imports and exports using the dynamic OLS methodology proposed by Stock and Watson (2003). In the cases of UK, Spain, Italy and USA, some of the estimated equations were not consistent with theory and thus we have proceeded without imposing long-run restrictions.

| Table 4.3: Lag order of individual VARX* model and number of cointegrating relations | | | |
|---|---|---|---|
| Countries | Lag order of domestic variables $p_j$ | Lag order of foreign variables $q_j$ | Number of cointegrating vectors |
| UK | 1 | 1 | 2 |
| GE | 1 | 1 | 1 |
| FR | 1 | 1 | 1 |
| SP | 1 | 1 | 2 |
| IT | 1 | 1 | 1 |
| US | 1 | 1 | 2 |

| Table 4.4: Cointegration results | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Trace statistic | | | | | | Maximum eigenvalue statistic | | | | | |
| | UK | GE | FR | SP | IT | US | UK | GE | FR | SP | IT | US |
| r=0 | 119,66 | 113,35 | 114,27 | 255,39 | 102,62 | 144,59 | 42,5 | 49,73 | 53,68 | 149,22 | 42,16 | 57,19 |
| r=1 | 77,16 | 63,61 | 60,58 | 106,16 | 60,45 | 87,40 | 38,89 | 37,18 | 28,11 | 64,56 | 22,46 | 52,67 |
| r=2 | 38,26 | 26,43 | 32,46 | 41,60 | 37,98 | 34,73 | 28,29 | 18,37 | 19,95 | 27,11 | 19,99 | 18,36 |
| r=3 | 9,97 | 8,05 | 12,51 | 14,48 | 17,99 | 16,36 | 9,97 | 8,05 | 12,51 | 14,48 | 17,99 | 16,36 |

**Notes:** Sample size: 73. The critical values for the trace statistic at the 5% significance level are: 100,96 (r=0), 71,56(r=1), 45,9(r=2), 23,63(r=3), MacKinnon, Haug, Michelis (1999).

### 4.4.5. Weak Exogeneity and structural stability tests

The VARX*(1,1) model can be written as a vector autoregressive error correction model in domestic and foreign variables, denoted as VECMX(1,1). Table 4.5 reports the *F*-statistics for serial correlation of the residuals of this model. On inspection we can see that for each country-specific model, the null hypothesis that the residuals are not serially correlated at 4 lags cannot be rejected at 5% significance level for all countries and all variables with the exception of real income in the case of Italy and the real effective exchange rate in the case of Spain.

**Table 4.5: Test for residual serial correlation of the VECMX***

| | | CV 5% | *z* | *x* | *y* | *e* |
|---|---|---|---|---|---|---|
| UK | F(4,58) | 2,530 | 2,295 | 0,304 | 2,014 | 1,740 |
| GE | F(4,59) | 2,527 | 0,160 | 1,914 | 0,916 | 1,346 |
| FR | F(4,59) | 2,527 | 1,513 | 0,933 | 0,506 | 2,284 |
| SP | F(4,58) | 2,530 | 1,018 | 1,018 | 1,328 | 15,127 |
| IT | F(4,59) | 2,527 | 0,560 | 1,540 | 2,588 | 2,217 |
| US | F(4,58) | 2,530 | 0,131 | 1,344 | 0,242 | 0,979 |

**Notes:** Sample size: 73. CV 5% is the critical value of the *F*-statistic at 5% significance level.

One of the main assumptions of the VARX*(1,1) model is that the foreign variables are weakly exogenous with respect to cointegrating relations. In Table 4.6 which reports the results we can see that the hypothesis of weak exogeneity cannot be rejected at 5% significance level for all countries.

| Table 4.6: Test for weak exogeneity | | | | | | |
|---|---|---|---|---|---|---|
| | | CV 5% | *z** | *x** | *y** | *e** |
| UK | F(2,57) | 3,158 | 0,106 | 0,180 | 0,292 | 0,005 |
| GE | F(1,58) | 4,006 | 0,219 | 1,021 | 0,045 | 0,634 |
| FR | F(1,58) | 4,006 | 0,177 | 0,002 | 0,0006 | 0,307 |
| SP | F(2,57) | 3,158 | 0,380 | 0,002 | 0,692 | 0,095 |
| IT | F(1,58) | 4,006 | 0,367 | 0,383 | 0,002 | 1,014 |
| US | F(2,57) | 3,158 | 0,359 | 0,412 | 0,006 | 0,628 |
| **Notes:** Sample size: 73. CV 5% is the critical value of the *F*-statistic at 5% significance level. | | | | | | |

Another important assumption of the VARX*(1,1) model is that the short-run parameters of each country-specific model do not exhibit structural breaks. The structural stability tests we have performed are the following:

- CUSUM statistic based on OLS residuals, denoted as PKsup, and its mean square variant, PKmsq (Ploberger and Kramer, 1992).
- Nyblom (1989) test for parameter constancy against non-stationary alternatives, and its heteroskedasticity-robust version, Robust Nyblom.
- Quandt’s (1960) likelihood ratio statistic (QLR) of a one-time structural change at an unknown change point, and its heteroskedasticity-robust version, Robust QLR.
- Hansen (1992), and Andrews and Ploberger (1994) mean Wald statistic (MW) of a one-time structural change at an unknown change point, and its heteroskedasticity-robust version, Robust MW.
- Andrews and Ploberger (1994) exponential average Wald statistic (APW) of a one-time structural change at an unknown change point, and its heteroskedasticity-robust version, Robust APW.

Table 4.7 presents the results. On inspection we observe that there is a broad evidence of parameter stability. The observed instability seems to result from breaks in the error variance and not the estimated coefficients, and once these breaks were allowed for by conducting the robust versions of the stability tests, the number of rejections of the null hypothesis has been reduced, indicating that the parameter coefficients seem to be reasonably stable. This evidence seems to justify why the likely break from the Global Financial Crisis is not taken into account. Following Dees et al. (2007), we have dealt with the possibility of breaks in the error variance by using bootstrap median estimates and confidence interval, rather than point estimates, of the impulse response functions. In addition, in the analysis of the short run effects, we report *t*-statistics which are based on robust standard errors.

**Table 4.7: Structural stability tests**

| | Number (percentage) | | | | |
|---|---|---|---|---|---|
| **Tests** | ***z*** | ***x*** | ***y*** | ***e*** | Total |
| PK sup | 1(16.7) | 0(0) | 1(16.7) | 3(50) | 5(20.8) |
| PK msq | 0(0) | 0(0) | 0(0) | 2(33.3) | 2(8.3) |
| Nyblom | 1(16.7) | 0(0) | 0(0) | 0(0) | 1(4.2) |
| Robust Nyblom | 1(16.7) | 0(0) | 0(0) | 0(0) | 1(4.2) |
| QLR | 2(33.3) | 1(16.7) | 2(33.3) | 2(33.3) | 7(29.2) |
| Robust QLR | 1(16.7) | 1(16.7) | 0(0) | 0(0) | 2(8.3) |
| MW | 1(16.7) | 0(0) | 2(33.3) | 1(16.7) | 4(16.7) |
| Robust MW | 1(16.7) | 1(16.7) | 0(0) | 0(0) | 2(8.3) |
| APW | 0(0) | 1(16.7) | 2(33.3) | 1(16.7) | 4(16.7) |
| Robust APW | 1(16.7) | 0(0) | 0(0) | 1(16.7) | 2(8.3) |

**Notes**: Sample size: 73. The entries are the number and the perecentage of rejections of the null hypothesis of parameter stability for each variable and country specific model at 5% significance level.

### 4.4.6. Short-run effects

The statistically significant contemporaneous effects of the foreign variables on domestic counterparts, which are interpreted as impact elasticities between these variables, are reported in Table 4.8. The short-run linkages between the real output growth rates are positive and very strong in UK, and moderate in France, Spain and Italy. In the UK, an increase of 1% in foreign output will lead to an increase be 1% in domestic output, thus implying a strong comovements between domestic and foreign output growth rates. The evidence also suggests that an increase in foreign imports will increase domestic imports in all countries except for Germany, and an increase in foreign exports will increase domestic exports in Germany, France and USA. The link between domestic and foreign trade seems to indicate that domestic exports and imports co-vary together over the business cycle, because of the strong import component of exports due to globalization and internationalization of production. Finally, an appreciation of the foreign real exchange rates will appreciate the domestic real exchange rate in the four euro area countries and will depreciate the real exchange rate of the dollar.

**Table 4.8: Short-run effects of foreign variables on domestic variables**

| | UK | GE | FR | SP | IT | US |
|---|---|---|---|---|---|---|
| ***z*** | 0,369** [2,14] | | 0,203** [1,90] | 0,441** [2,53] | 0,955** [4,41] | 0,359** [2,04] |
| ***x*** | | 0,502** [2,07] | 0,358** [2,22] | | | 0,65** [4,46] |
| ***y*** | 0,985** [3,98] | | 0,337** [3,28] | 0,683** [5,98] | 0,566** [4,55] | |
| ***e*** | | 0,552** [2,79] | 1,160** [10,72] | 1.131** [5,32] | 1,276** [11,49] | -0,972** [-4,00] |

**Notes:** Sample size: 73. Numbers in square brackets are White's *t*-ratios. ** indicates significance at 5% level.

### 4.4.7. Pair-wise cross-section correlations

We have also examined whether the idiosyncratic shocks of the individual countries are weakly correlated, implying that the covariance between the foreign variables and the error terms of the VARX* model are zero, as the number of countries in the model approaches infinity and thus the weak exogeneity of foreign variables of the model is established. In Table 4.9, we present the average pair wise cross-section correlations of the levels and first-differences of variables, and the estimated residuals of the VECMX*. The results indicate that the cross-section correlations are higher for the levels, with a few exceptions, than for the first differences, and very small for the residuals. The levels of exports show the greater degree of cross-section correlations, ranging from 90% to 96%, and this is followed by imports, ranging from 70% to 88%, real output, ranging from 66% to 72%, and finally exchange rates, ranging from -10% to 58%. The growth rates of real output show the greater degree of cross-section correlations, ranging from 51% to 68%, and followed by the growth rates of imports and exports, ranging from 43% to 63%, and 42% to 66%, respectively. Finally, the growth rate of real exchange rates ranges from -45% to 39%.

With respect of the cross-correlations of the residuals of the VECMX*, the evidence suggests that these correlations are quite smaller than those obtained from the levels and the first differences. In particular, out of 24 computed correlations, 5 are in the range 0 to 15%, 12 are in the range -10% to 0, 5 are in the range -20% to -11%, and two have correlations -28% and -36%.

Overall, the analysis indicates that the variables of the GVAR model exhibit significant cross-country correlations for both the levels and the first differences, whereas there exist a modest degree of correlations across shocks from different countries.

| Table 4.9: Average pair wise cross-section correlations | | | | |
|---|---|---|---|---|
| | | Levels | First Differences | VECMX* Residuals |
| ***z*** | UK | 0,880 | 0,429 | -0,057 |
| ***z*** | GE | 0,843 | 0,529 | -0,146 |
| ***z*** | FR | 0,855 | 0,591 | -0,070 |
| ***z*** | SP | 0,708 | 0,585 | -0,030 |
| ***z*** | IT | 0,790 | 0,598 | 0,001 |
| ***z*** | US | 0,883 | 0,632 | -0,040 |
| ***x*** | UK | 0,945 | 0,416 | -0,132 |
| ***x*** | GE | 0,960 | 0,649 | -0,164 |
| ***x*** | FR | 0,963 | 0,653 | -0,127 |
| ***x*** | SP | 0,956 | 0,589 | -0,083 |
| ***x*** | IT | 0,906 | 0,667 | -0,097 |
| ***x*** | US | 0,934 | 0,628 | -0,069 |
| ***y*** | UK | 0,719 | 0,601 | -0,029 |
| ***y*** | GE | 0,663 | 0,587 | -0,091 |
| ***y*** | FR | 0,724 | 0,656 | -0,129 |
| ***y*** | SP | 0,692 | 0,533 | -0,015 |
| ***y*** | IT | -0,080 | 0,675 | -0,022 |
| ***y*** | US | 0,706 | 0,517 | -0,095 |
| ***e*** | UK | 0,281 | -0,075 | -0,283 |
| ***e*** | GE | 0,582 | 0,377 | 0,150 |
| ***e*** | FR | 0,568 | 0,390 | 0,076 |
| ***e*** | SP | 0,150 | 0,296 | 0,001 |
| ***e*** | IT | 0,497 | 0,380 | 0,047 |
| ***e*** | US | -0,103 | -0,451 | -0,358 |

### 4.4.8. Impulse response analysis

We use the estimated GVAR model to perform impulse response analysis of the the effects that shocks to selected variables of the model have on the economy of all countries in the analysis. In particular, we use generalized impulse response functions, which are invariant to the ordering of the variables in the system, to simulate a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area. An interesting aspect of the impulse response analysis is the evidence that both shocks spillover to all the countries in the analysis, indicating that the trade relations between them constitute the transmission mechanism of real shocks between them.

*A negative shock to UK GDP*

Figures 4.2 plots the response of the real output over a horizon of 40 quarters to one standard error negative shock to UK real output. The numbers refer to bootstrap median estimates with 90% bootstrap error bounds. The negative shock is transmitted quickly to other countries and is accompanied by a decline in real output in all countries during the first year. In particular, it takes about two quarters for real output to respond significantly to the negative shock. Then, it remains pronounced and statistically significant for the entire forecasting horizon. The one-standard deviation negative shock means a fall of UK GDP by 0,38% at the time of impact. After one year, the GDP in Germany has decreased by 0,67% which is the largest impact, indicating the strong trade interactions between the two countries, in Italy has fallen by 0,54%, in Spain 0,39%, and France 0,36%. Interestingly, we observe that the negative shock to UK GDP also reduces the GDP in USA by 0,32%. Given the fact that the USA is the largest destination of UK exports and the second largest origin of UK imports, this finding seems to suggest that an adverse real output shock affects the USA through trade relations (see Figure 4.3). A recent study (Bodenstein et al. 2009) presents evidence that an adverse foreign demand shock which reduces foreign real output by 1% will cause a reduction in the USA real output by 0.3% if interest rates response to this shock. The reduction is larger when the USA is in a liquidity trap and interest rates are constrained to a low zero bound. If the foreign policy rate is reduced to stabilize the economy, then the associated depreciation of the foreign currency would cause a reduction in USA exports, in addition to the reduction resulting from the fall in foreign demand. On the other hand, if the USA policy rate does not respond, then the fall in expected inflation will increase short-term real interest rates and consequently real demand and output.

Figures 4.3 plots the responses of real imports and exports to one standard error negative shock to UK real output. The numbers refer to bootstrap median estimates with 90% bootstrap error bounds. The negative shock reduces imports and exports in all countries, but the impact is different across variables and countries, implying a different response of trade balances. During the first year, for instance, the trade deficit in the UK increases as imports decline 1,4% and exports 1,5%, but decreases in the USA as imports fall 1,68% and exports fall 1,47%. On the other hand, the trade surplus in Germany decreases as exports fall 2,04% and imports fall 1,39%. Interestingly, the lower output growth in the UK has a negative effect on foreign exports implying a lower output growth in other countries something which discussed before.

Figure 4.2: Generalized impulse responses of real output to one standard error negative shock to UK real output (bootstrap median estimates with 90% bootstrap error bounds)

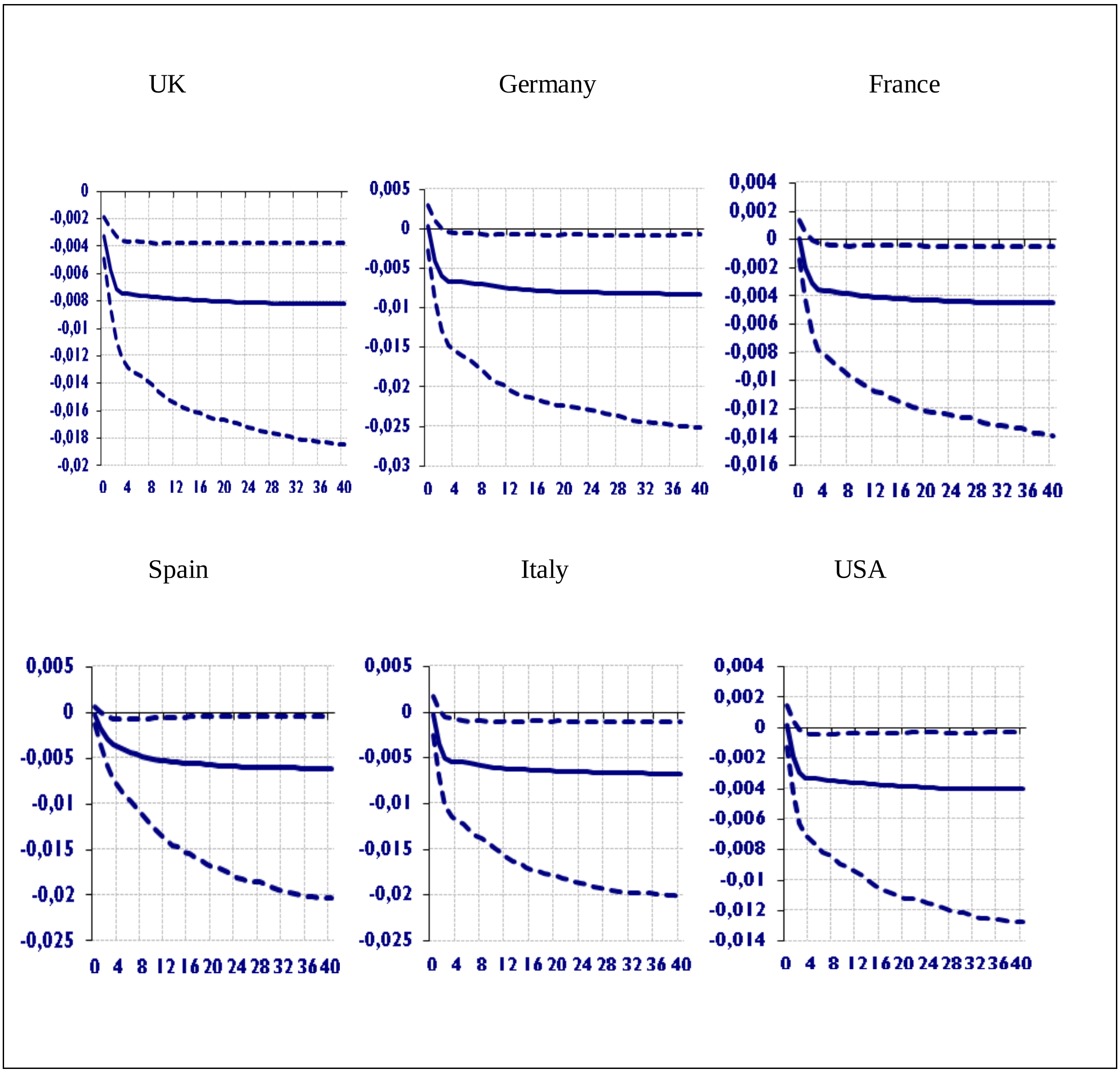
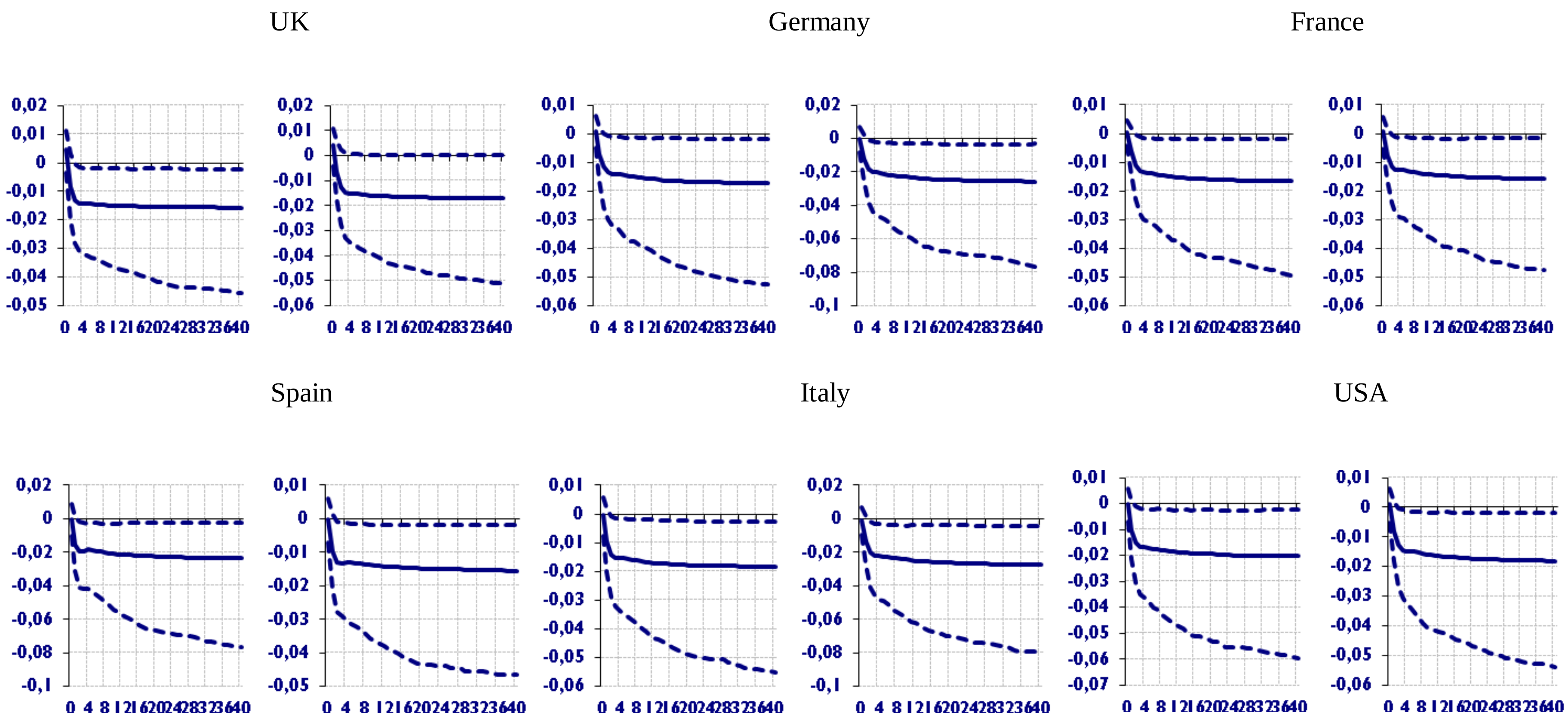


Figure 4.3: Generalized impulse responses of real imports and real exports to one standard error negative shock to UK real output (bootstrap median estimates with 90% bootstrap error bounds)



*A positive shock to German GDP*

Figures 4.4 plots the response of the real output over a horizon of 40 quarters to one standard error positive shock to German real output. The numbers refer to bootstrap median estimates with 90% bootstrap error bounds. The positive shock is transmitted quickly to other countries and is accompanied by an increase in real output in all countries during the first year. Then, it remains pronounced and statistically significant for the forecasting horizon. The one-standard deviation positive shock means a rise of German GDP by 0,5% at the time of impact. After one year, the GDP in the UK has increased 0,59%, in Italy 0,77%, in Spain 0,54% and in France 0,52%. Interestingly, we observe that the positive shock to German real output also increases the real output in USA by 0,44%. The higher output growth in Germany has a positive effect on foreign exports implying a higher output growth in other countries. This is evident from Figure 4.5 which plots the responses of real imports and real exports to one standard error postive shock to German real output. The numbers refer to bootstrap median estimates with 90% bootstrap error bounds. The positive shock increases imports and exports in all countries, but the impact is different across variables and countries, implying a different response of trade balances. During the first year, for instance, the trade deficit in the UK falls as imports increase 1,4% and exports 1,83%, but rises in the USA as imports increase 2,42% and exports 1,95%. On the other hand, the trade surplus in Germany increases as exports increase 3,52% and imports 2,36%.

Figure 4.4: Generalized impulse responses of real output to one standard error positive shock to German real output (bootstrap median estimates with 90% bootstrap error bounds)

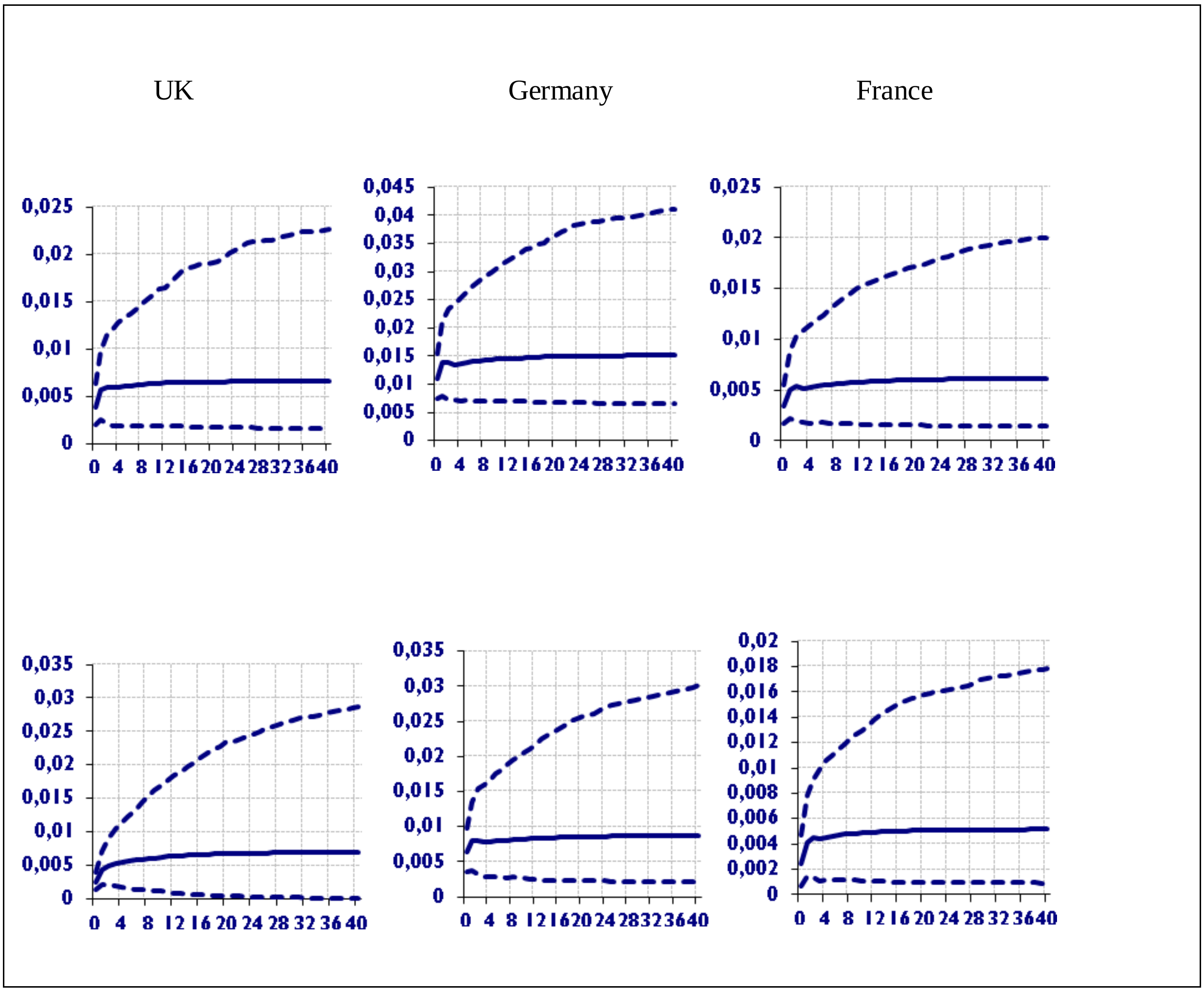
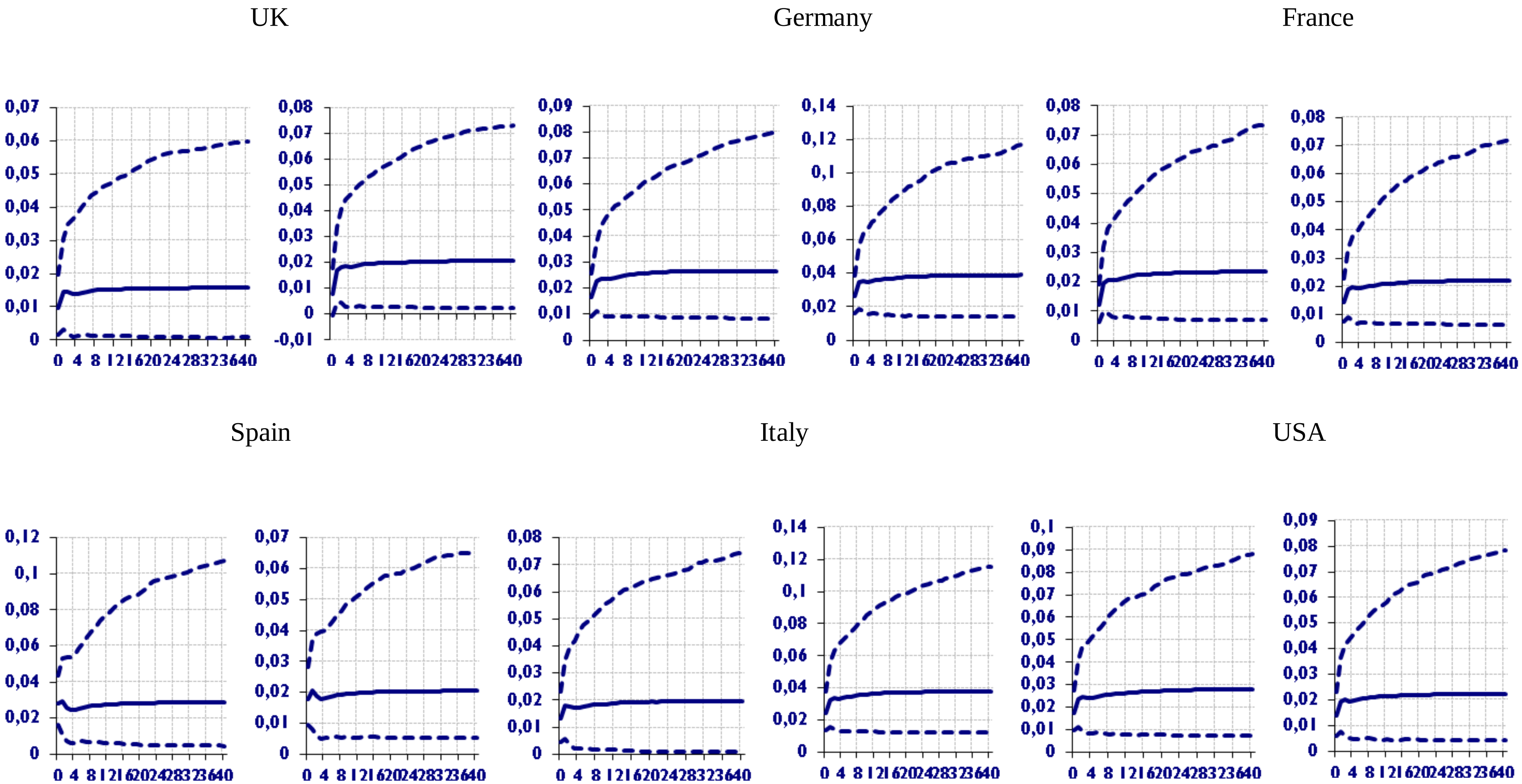


Figure 4.5: Generalized impulse responses of real imports and real exports to one standard error positive shock to German real output (bootstrap median estimates with 90% bootstrap error bounds)



### 4.4.9. Forecast error variance decomposition

We use the estimated GVAR model to perform forecast error variance decomposition of the the effects that shocks have on selected variables of the model. In particular, we use generalized forecast error variance decomposition, which are invariant to the ordering of the variables in the system, to compute the proportion of the 12-quarter ahead forecast errors of imports and exports of the UK and Germany that is explained by conditioning on current and future values of shocks to all variables of the system. The analysis shows that domestic and foreign shocks contribute to the forecast variance of real imports and real exports in the two countries, and that traditional trade equations, which relate real imports to domestic real income and the exchange rate, and real exports to foreign real income and the exchange rate, are not supported by the data.

*Real imports and exports of the UK*

Figure 4.6 plots the results for those variables which explain a proportion of 10% or more of the forecast variance of real imports at the twelve-quarter horizon. In one-year horizon, the domestic real income and the real exchange rate explain 23% of the forecast variance of real imports which is less than a quarter of the total variance. In a three-year horizon, the domestic real income and the real exchange rate explain 27% of the forecast variance which amounts to 20% of the total variance. At both horizons, the contribution of the real exchange rate is more important. However, the small contribution of both variables indicates that real imports are not primarily affected by the traditional determinants. At the both horizon, real imports and real exports explain 24% of the forecast variance of real imports, and thus all domestic variables together account for a half proportion of the total variance of real imports. The other half proportion of the total variance is explained by foreign variables, such as German real imports, real exports and real income, and USA real exports.

Figure 4.7 plots the results for those variables which explain a proportion of 10% or more of the forecast variance of real exports at the twelve-quarter horizon. In one-year horizon, the foreign (German) real income and the real exchange rate explain 23% of the forecast variance of real exports which is less than a quarter of the total variance. In a three-year horizon, the foreing real income and the real exchange rate explain 25% of the forecast variance, amounting to a quarter of the total variance. At both horizons, the contribution of the foreign income is more important. However, the small contribution of both variables indicates that real exportsts are not primarily affected by the traditional determinants. At the both horizon, real imports and real exports explain 35% of the forecast variance of real exports, and thus all domestic variables together account for a half proportion of the total variance of real exports. The other half proportion of the total variance is explained by foreign variables, such as German real imports, real exports and real income, and USA real exports.

Figure 4.6: Generalised forecast error variance decomposition of UK real imports

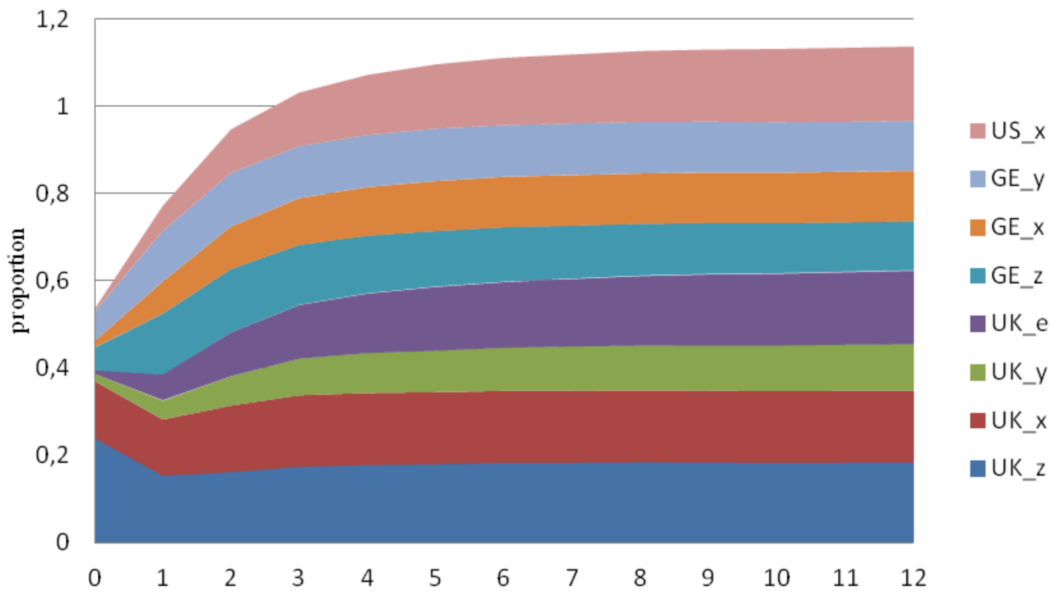
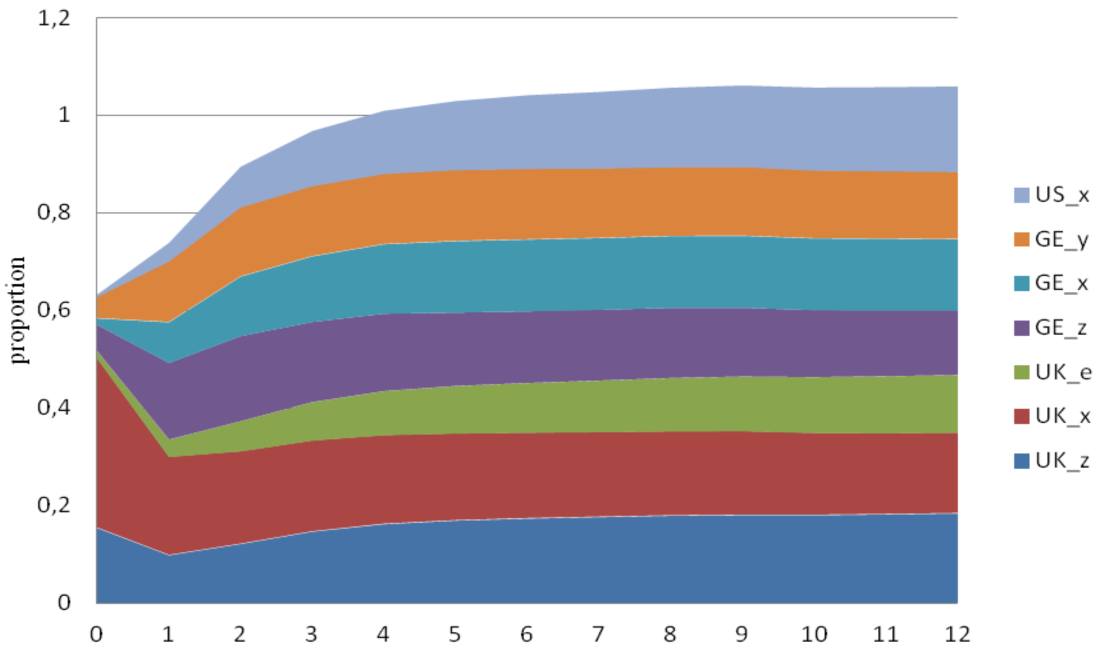



Figure 4.7: Generalised forecast error variance decomposition of UK real exports



*Real imports and exports of Germany*

Figure 4.8 plots the results for those variables which explain a proportion of 10% or more of the forecast variance of real imports at the twelve-quarter horizon. In one-year horizon, the domestic real income explains 24% of the forecast variance of real imports which is about 20% of the total variance. In a three-year horizon, the domestic real income explain 20% of the forecast variance, amounting to 17% of the total variance. At both horizons, the contribution of the real exchange rate is 2%. These results indicate that real imports are not primarily affected by the traditional determinants. At the short horizon, real imports and real exports explain 45%, and at the long horizon 40%, and thus all domestic variables together account for two-third of the total variance of real imports. The proportion of the total variance is explained by foreign variables, such as UK real imports, real exports, and USA real exports.

Figure 4.9 plots the results for those variables which explain a proportion of 10% or more of the forecast variance of real exports at the twelve-quarter horizon. At both horizons, neither the foreign income nor the real exchange rate affect at all the real exports, suggesting that real exports are not affected by the traditional determinants. In one-year horizon, domestic variables, that is real imports, real exports, and real income, explain 73% of the forecast variance of real exports, amounting to 57% of the total variance. In a three-year horizon, domestic variables explain 62% of the forecast variance of real exports, amounting to 50% of the total variance. The other half proportion of the total variance is explained by foreign variables, such as UK real imports, and real exports, and USA real exports.

Figure 4.8: Generalised forecast error variance decomposition of German real imports

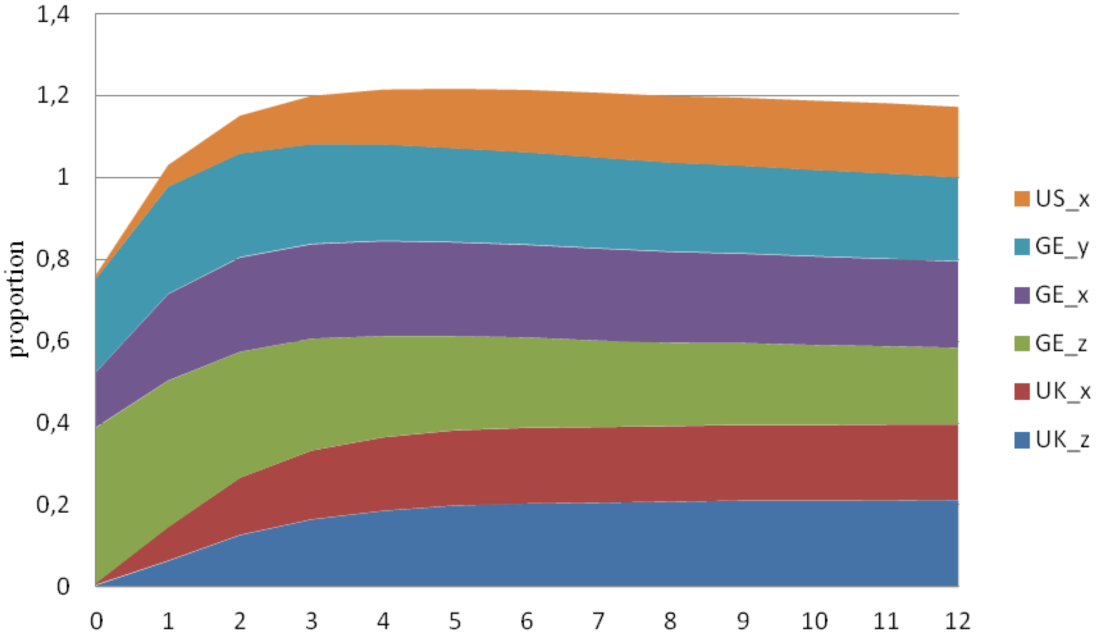
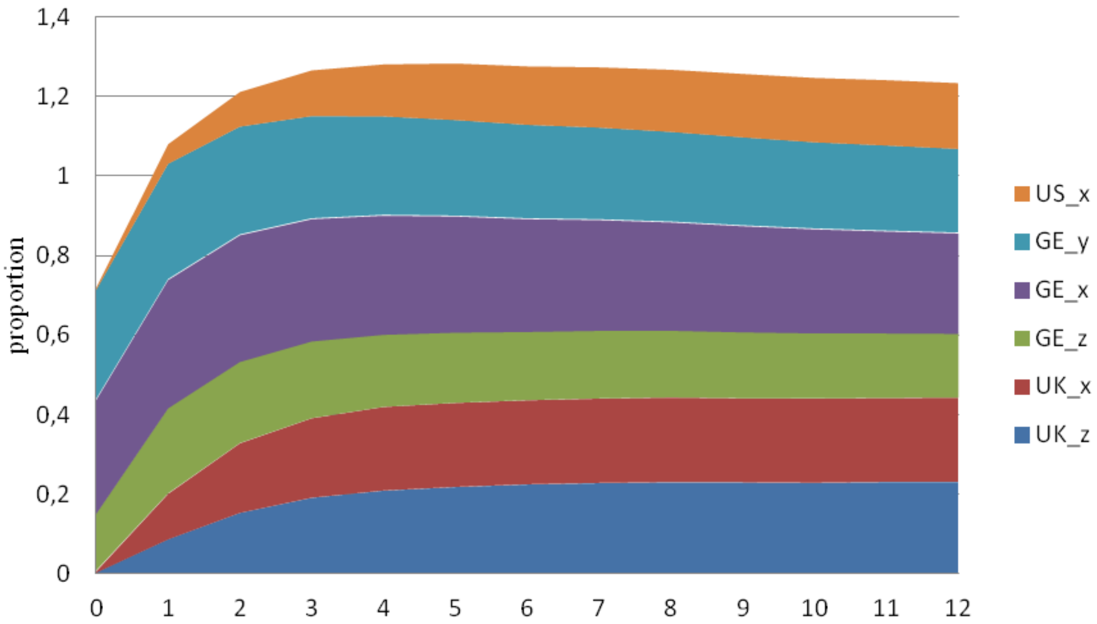



Figure 4.9: Generalised forecast error variance decomposition of German real exports



## 4.5. Conclusion

The outcome of Britain's referendum to leave the EU has spurred a discussion about the potential risks of this decision on the UK economy, since the EU constitutes its largest trading partner. In the empirical literature, there exist a few papers which examine the potential effects of Brexit. The main conclusion which results from these studies is that these effects would be substantial.

This chapter contributes to the empirical literature on the potential risks of Brexit by using an alternative approach to model the trade interactions between the UK and the EU, and forecast whether these historical trade flows provide information about the impact of Brexit on the UK economy. The alternative approach we use to answer the research question is the innovative tool of GVAR. In particular, we use the model to forecast the effects of a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area, on the economy of all countries in the analysis. In addition, we compute the proportion of the forecast variance of real imports and real exposts of the UK and Germany that is explained by shocks to all variables and countries in the model.

Initially, we have constructed the foreign variables based on the figures of the weight matrix. These figures were calculated as the ratio of the bilateral trade flows between the six countries to the total trade of each country during the period 2012-2017. The figures indicate that there exist significant trade interrelations between the countries which play key role in the transmission of real shocks between them. The cointegration analysis has shown that for each specific country there exist a different number of cointegrating relationships among the four variables. In particular, two cointegrating relations exist for UK, Spain and USA, and one cointegrating relation for the other three countries. Then, we have used the dynamic OLS methodology to examine whether these cointegrating relations are consistent with theory. The analysis shows that some of the estimated equations were not consistent with theory in terms of the significance and the expected signs of the estimated coefficients, and thus we have proceeded without imposing long-run restrictions.

In the explicit model for each country, we have tested the main assumption that the foreign variables are weakly exogenous with respect to cointegrating relations. The hypothesis of weak exogeneity cannot be rejected for all countries in the analysis. Another important assumption which has been examined is that the short-run parameters of each country-specific model do not exhibit structural breaks. In general, we observe that there is a broad evidence of parameter stability. The observed instability seems to result from breaks in the error variance and not the estimated coefficients, and once these breaks were allowed for by conducting the robust versions of the stability tests, the number of rejections of the null hypothesis has been reduced, indicating that the parameter coefficients seem to be reasonably stable. This evidence seems to justify that the likely break from the Global Financial Crisis is not taken into account. We have dealt with the possibility of breaks in the error variance by using bootstrap median estimates and confidence interval, rather than point estimates, of the impulse response functions. In addition, we use robust standard errors in the analysis of the short-run effects. This short-run analysis, shows strong contemporaneous linkages between the variables of the system. In particular, we observe that the real output growth rates are positive and very strong in UK, and moderate in France, Spain and Italy. In the UK, an increase of 1% in foreign output will lead to an increase be 1% in domestic output, thus implying a strong comovements between domestic and foreign output growth rates. The evidence also suggests that an increase in foreign imports will increase domestic imports in all countries except for Germany, and an increase in foreign exports will increase exports in Germany, France and USA. The link between domestic and foreign trade seems to indicate that domestic exports and imports co-vary together over the business cycle, because of the strong import component of exports due to globalization and internationalization of production. Finally, an appreciation of the foreign real exchange rates will appreciate the domestic real exchange rate in the four euro area countries and will depreciate the real exchange rate of the dollar.

We have used generalized impulse response functions, which are invariant to the ordering of the variables in the system, to simulate a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area. The negative shock is transmitted quickly to other countries and is accompanied by a decline in real output in all countries during the first year. In particular, it takes about two quarters for real output to respond significantly to the negative shock. Then, it remains pronounced and statistically significant for the entire forecasting horizon. The one-standard deviation negative shock means a fall of UK GDP by 0,38% at the time of impact. After one year, the GDP in Germany has decreased by 0,67% which is the largest impact, indicating the strong trade interactions between the two countries. Interestingly, the negative shock also reduces the GDP in USA by 0,32%, suggesting that the adverse real output shock affects the USA through the trade channel, given the evidence that the lower output growth in the UK has a negative effect on the US real exports. In addition, the negative shock reduces real imports and real exports in all countries, but the impact is different across variables and countries, implying a different response of trade balances. The positive shock to German real output is also transmitted quickly to other countries and is accompanied by an increase in real output in all countries during the first year. Then, it remains pronounced and statistically significant for the entire forecasting horizon. The one-standard deviation positive shock means a rise of German GDP by 0,5% at the time of impact. After one year, the GDP has increased 0,59% in the UK and 0,44% in the USA. The higher output growth in Germany has a positive effect on foreign real exports implying a higher output growth in other countries. In addition, the impact on real imports and exports is different across variables and countries, implying a different response of trade balances.

Finally, we have used generalized forecast error variance decomposition to compute the proportion of the 12-quarter ahead forecast errors of imports and exports of the UK and Germany that is explained by conditioning on current and future values of shocks to all variables of the system. We have reported the results for those variables which explain a proportion of 10% or more of the forecast variance of real imports and real exports at the twelve-quarter horizon. In the UK, the domestic real income and the real exchange rate explain 23% of the forecast variance of real imports in one-year horizon, which is less than a quarter of the total variance, and 27% of the forecast variance of real imports in a three-year horizon, amounting to 20% of the total variance. At both horizons, the contribution of the real exchange rate is more important. However, the small contribution of both variables indicates that real imports are not primarily affected by the traditional determinants. In Germany, on the other hand, the domestic real income explains 24% of the forecast variance of real imports in one-year horizon, which is about 20% of the total variance, and 20% of the forecast variance of real imports in a three-year horizon, amounting to 17% of the total variance. At both horizons, the contribution of the real exchange rate is negligible. Also, at both horizons, neither the foreign income nor the real exchange rate affect at all the real exports, and thus real exports are not primarily affected by the traditional determinants.

In a nutshell, the conclusion which results from the GVAR analysis and answer the research question is that shocks to real output in the UK and Germany are significant and long-lasting for all economies in the analysis, and operate through the trade channel, and thus these results fill in the gap in the literature about the potential risks associated with Brexit.

# Chapter 5

# Conclusions

This dissertation discusses three essays on macroeconomic issues of the UK economy. The first essay focused on the dynamic covariance between stock and house returns and how its pattern could be influenced by monetary policy variables. The importance of this topic stems from the fact that, firstly, stock and house markets constitute transmission mechanisms of monetary policy, and secondly, stocks and houses constitute large components of the total wealth of UK households, thus affecting consumption and investment decisions. The discussion of the empirical literature reveals the gap referring to the co-movements of stock and house returns and the role of monetary policy developments in the UK in driving changes to the dynamic correlations of these returns. Given that asset price co-movements are of great importance to policy makers and portfolio investors, this chapter aspires to fill this gap. For this research, the DCC model was firstly implemented for the estimation of the covariance between the stock and house returns, and then a least squares with breaks method was used to filter the variability of dynamic conditional covariance between the stock market return and house market return. The variability constitutes the cyclical component of dynamic covariance which reflects periods of positive and negative deviations from trend. If deviations are large, boom and bust outcomes are present and imply risk. Managing risk means having information about the whole distribution of these possible outcomes. To get information about these outcomes, we have related the variability of dynamic conditional covariance with developments in monetary policy. Particularly, we have examined whether predetermined monetary policy contain information for forecasting changes in the dynamic covariance of stock-house returns across the entire conditional distribution of the response variable. To this end, we have used a quantile approach that is more informative than least squares which looks only at the conditional mean. The analysis shows that movements in M0 and M4 and the official interest rate are important to understanding future movements in the variability of dynamic correlations between the stock market return and house market return, given the information content of other predictors and variability's own past history. In particular, the three-month and six-month moving averages of both measures have a statistically significant impact on current variability of dynamic correlations, and the sign of their impact is positive in the higher quantiles and negative in the lower quantiles. This asymmetric effect implies that when changes in dynamic correlations are positive, an increase in the money supply will increase the variability of correlations of stock-house returns, thus increasing the possibility of a boom episode, whereas when changes in dynamic correlations are negative, an increase in the money supply will decrease the variability of stock and house returns, thus increasing the possibility of a bust episode. On the other hand, the three-month and six-month moving averages of changes in the official rate have a statistically significant impact on changes in dynamic correlations, and the sign of its impact is negative in the higher quantiles and positive in the lower quantiles. This asymmetric interest rate effect implies that when variability is positive, a fall in the official rate will further increase the variability of stock-house returns, thus increasing the possibility of a boom episode, whereas when variability is negative, a fall in the interest rate will decrease the variability of stock-house returns, thus increasing the possibility of a bust episode.

The conclusion which results from the quantile analysis and answers the research question is that developments in M0, M4 and the official interest rate have characterized the stylized facts of dynamic covariance between stock and house returns during the sample period, and hence provide valuable predictive content for the variability of the dynamic covariance. This evidence from the UK adds to the existing literature that stock and house markets constitute mechanisms though which monetary policy decisions affect the macroeconomic environment, and thus these facts should be taken into consideration when policy decisions are designed.

The second essay worked on the relationship between real output growth and debt-to-GDP ratio in the UK over the last three centuries. The main conclusion which results from the relevant literature is that the empirical studies have examined the link between public debt and real output growth without capturing an empirical fact that the growth rates of real GDP exhibit a more persistent profile during expansion periods than contraction periods. This empirical regularity cannot be explained by linear models used in the literature which cannot capture these distinct facts, but instead a nonlinear Markov-switching model capturing these distinct patterns in the data is regarded more appropriate. Thus, this chapter aspires to fill the gap in the literature referring to a non-linear analysis of the asymmetric behavior over real output expansions and contractions, and the role of the debt-to-GDP ratio in driving the transition probabilities between these dynamic structures. In particular, this chapter adds to the literature by examining the information content of the growth rate of the debt-to-GDP ratio in understanding the asymmetric behavior over real output expansions and contractions, using annual data from the middle of eighteenth century until 2016. In addition, the chapter examines whether the growth rate of the debt-to-GDP ratio is a leading indicator of the time-varying transition probabilities between expansions and contractions of the real output growth. The importance of this topic stems from the fact that the link between debt and output growth is crucial from a policy perspective, as the establishment of negative growth effects of government debt will support the arguments for long-run fiscal consolidation.

We have specified a backward-looking aggregate demand equation which relates the growth rate of real output on its own past value, the past value of the debt-GDP ratio, and past values of the change in the real long-term interest rate, the inflation rate, and the log first difference of the real exchange rate. The model allows for different sources of regime shifts, such as changes in the intercept, the dynamics of real output, the debt-to GDP ratio, and the error variance. In this context, the two regimes are identified as high and low growth rates of real output. Initially, we have estimated the model with constant transition probabilities under the assumption that the mean growth rate of real output, the propagation mechanism, the lagged value of the debt-to-GDP ratio, and the error variance are subject to regime shifts. The results suggest that the presence of nonlinearities is related to different means, volatilities and dynamics of real output growth across samples, and the lagged value of the growth rate of the debt-to GDP ratio has predictive content useful to understand movements in the growth rate of the real output in the high growth regime. The negative sign shows that an increase in the public debt relative to GDP is expected to decrease the growth rate of real output. Then, we have estimated the model with time-varying transmission probabilities, which depend on the lagged value of the growth rate debt-to-GDP ratio. The results with common and error specific variances showed that increases in the lagged growth rate of the debt ratio are associated with a higher probability of moving from the high growth regime to low growth regime. Thus, the debt ratio constitutes a leading indicator of the transition probability across the two regimes. The implication of these findings which fill in the gap in the literature on the link between real output growth and public debt support the arguments for long-run fiscal consolidation.

The third essay examined the trade relations between the UK and the Eurozone economies as a way of understanding the potential effects of Brexit on the economy of all countries in the analysis. Although some studies suggest that the potential costs of Brexit to the UK economy could be substantial, the literature is rather limited, and this requires more research to examine the significance and duration of the potential risks for the UK and the EU economies. This chapter aspires to fill in this gap in the literature by using the innovative tool of the global vector autoregression (GVAR) to model the trade interactions between the UK, the four big euro area countries – Germany, France, Italy and Spain – and the USA. Then, we use the model to perform impulse response analysis and forecast error variance decomposition. In particular, we forecast the effects of a negative shock to UK real output, which proxies the uncertainty surrounding the withdrawal from the EU, and a positive shock to Germany real output, which proxies an expansionary shock in the euro area, on the economy of all countries in the analysis. In addition, we compute the proportion of the forecast variance of real imports and real exports of the UK and Germany that is explained by shocks to all variables and countries in the model.

The impulse response analysis shows that the negative shock is transmitted quickly to other countries and is accompanied by a decline in real output in all countries during the first year. Then, it remains pronounced and statistically significant for the entire forecasting horizon. The one-standard deviation negative shock, which means a fall of UK GDP by 0,38% at the time of impact, decreases the German GDP by 0,67%. It is worth noting that the adverse real output shock operates through the trade channel, as it reduces real imports and real exports in all countries. Interestingly, the positive shock to German real output is also transmitted quickly to other countries and is accompanied by an increase in real output in all countries during the first year. Then, it remains pronounced and statistically significant for the entire forecasting horizon. The one-standard deviation positive shock, which means a rise of German GDP by 0,5% at the time of impact, increases the UK GDP by 0,59% after one year. The expansionary shock in the first economy of the euro area operates through the trade channel, as it increases real imports and real exports in all countries.

The forecast error variance decomposition analysis shows that in the UK the domestic real income and the real exchange rate explain 23% of the forecast variance of real imports in one-year horizon, which is less than a quarter of the total variance, and 27% of the forecast variance of real imports in a three-year horizon, amounting to 20% of the total variance. At both horizons, the contribution of the real exchange rate is more important. However, the small contribution of both variables indicates that real imports are not primarily affected by the traditional determinants. In Germany, on the other hand, the domestic real income explains 24% of the forecast variance of real imports in one-year horizon, which is about 20% of the total variance, and 20% of the forecast variance of real imports in a three-year horizon, amounting to 17% of the total variance. At both horizons, the contribution of the real exchange rate is negligible. The very small contribution of the foreign income and the real exchange rate in explaining the forecast variance of the real exports implies that real exports are not primarily affected by the traditional determinants. The implications of these findings which fill in the gap in the literature about the potential risks associated with the Brexit are that shocks to real output in the UK and Germany are significant and long-lasting for all economies in the analysis, and operate through the trade channel.

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