Credit and Capital Markets, Volume 52, Issue 2, pp. 173–190
Scientific Papers

The Interest Rate Sensitivity of Investment

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Abstract

The interest rate sensitivity of investment has often played an important role in macroeconomic models. However, many vector autoregressive (VAR) models do not include investment to the list of variables. In this paper, we empirically investigate the size and the evolution of the interest rate sensitivity of investment for the United States and the four largest European economies in the last few decades. We use a VAR model with four variables at quarterly frequency: real investment, real gross domestic product (GDP), inflation, and a measure of the short-term interest rate. In our VAR, the structural interest rate shock is identified under the assumption that macroeconomic quantities and inflation react to interest rate innovations with a lag. We test the appropriateness of this specification by comparing our approach with the identification of shocks derived from the changes in volatility approach. For the countries under consideration, we determine a date during either the 1980s or the 1990s where the interest rate sensitivity of investment began to decrease and became less responsive to monetary policy. In addition, we find that the interest rate sensitivity of investment has been higher in the United States than in Europe, particularly in the first subperiod.

Die Zinssensitivität der Investitionen

Zusammenfassung

Die Zinssensitivität der Investitionen spielt oft eine große Rolle in theoretischen makroökonomischen Modellen. In dieser Studie untersuchen wir empirisch die Höhe und die zeitliche Änderung der Zinssensitivität der Investitionen für die Vereinigten Staaten und die vier größten europäischen Volkswirtschaften. Wir verwenden ein VAR-Modell

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We thank the anonymous referees and the editor Hendrik Hakenes for very useful comments and suggestions. In addition, comments by various colleagues at our institutions are gratefully acknowledged. All remaining errors are our own. The views expressed in this paper are those of the authors and not necessarily those of the institutions to which the authors are affiliated.

mit vier Variablen: reale Investitionen, reales Bruttoinlandsprodukt, Inflation und kurzfristige Zinsen. In unserem VAR identifizieren wir den strukturellen Schock unter der Annahme, dass die realen makroökonomischen Variablen verzögert auf einen Zinsschock reagieren. Wir testen die Angemessenheit dieser Spezifikation, indem wir unsere Vorgehensweise mit der Identifikation durch den "changes in volatility approach" vergleichen. Wir finden heraus, dass entweder in den 1980er oder frühen 1990er Jahren ein Strukturbruch stattgefunden und sich die Zinssensitivität der Investitionen verringert hat. Interessanterweise zeigen unsere Resultate zudem, dass die Zinssensitivität der Investitionen in den Vereinigten Staaten höher gewesen ist als in den untersuchten europäischen Ländern – insbesondere bis in die 1980er Jahre.

Keywords: Investment, Effects of Interest Rates, Monetary Policy

JEL Classification: E22, E43, E52

I. Introduction

The interest rate sensitivity of investment has played an important role in many macroeconomic models. In the standard New Keynesian model, the short-term interest rate targeted by the central bank transmits the effects of monetary policy to the economy (for an overview, see e.g. *Gali* 2010). Our aim in this paper is to empirically investigate the size of the interest rate sensitivity of investment and its evolution over the recent economic history for the United States and the four largest European economies. We use a vector autoregressive (VAR) model with four variables at quarterly frequency: real investment, real gross domestic product (GDP), inflation, and a measure of the short-term interest rate. In many VAR models analyzing the effects of interest rate shocks, investment is not included in the list of variables under investigation. In addition, to the best of our knowledge, potential structural breaks in the interest rate sensitivity of investment have not been analyzed in the literature. In our analyses, we focus on the United States and the four largest European economies and apply a common strategy to identify interest rate shocks for these countries. Using a common model allows us to obtain a general picture of the effects of interest rate shocks. Because we are interested only in the effects of the interest rate shock, our main goal is to identify the interest rate innovations without having to recover the other structural shocks.

In the related literature using VAR models, the assumption is often made that macroeconomic quantities such as GDP respond to an interest rate shock with a lag. This is a priori plausible given the various reaction and implementation lags that affect macroeconomic quantities and the price setting process. However, one may also question the appropriateness of this identification strategy.[1] Following *Lanne/Lütkepohl* (2008b), the changes in volatility approach allows us to

[1] For alternative approaches, see e.g., *Christiano/Eichenbaum/Evans* (1999).

identify all the structural shocks without further restrictions and to test this unrestricted model against our identification. The changes in volatility approach can be used if the volatilities of the shocks differ across sub-periods in our sample. This is a pattern that one may often find in historical time series. Using Chow tests, we find evidence for a structural break point for each country under investigation. These break points occur during either the 1980s or the early 1990s, when the interest rate sensitivity of investment started to decrease. Having verified our identification strategy using the break points determined by the Chow tests, we then split the sample into two sub-periods and find that the sensitivity has decreased since the 1980s and early 1990s. According to our results, expansionary interest rate shocks have had expansionary effects on real investment in the past. In recent decades, interest rate shocks have displayed ambiguous real effects on investment and may be neutral. In addition, our findings imply that the interest rate sensitivity of real investment in the first subperiod was higher in the United States than in Continental Europe. The difference is less pronounced in the second subperiod.

Discussions on the relation between interest rates and investment have a long tradition that dates back at least as far as *Wicksell* ([1898] 1936) and *Klein* (1947) (for an overview, see e.g., *Backhouse/Boianovsky* 2016). In theory, investment can be expected to be influenced by a rise in the user cost of capital, of which the interest rate is one component.[2] Therefore, a decrease in interest rates is expected to lead to an increase in investment. Despite these clear theoretical predictions, empirical papers report difficulties in determining the sensitivity of investment to the interest rate or, more generally, to the user cost of capital (see, e.g. *Guiso/Kashyap/Panetta/Terlizzese* 2002). This is due to a potential simultaneity problem, especially when annual data are used, which may be reflected in a positive correlation between interest rates and actual or expected investment. Intuitively, it occurs because more optimistic expectations typically increase both interest rates and the number of profitable investment opportunities. In addition, a central bank usually seeks to increase its targeted interest rate when it expects investment to rise. If this simultaneity problem is not addressed in empirical analyses, estimations of the interest rate sensitivity may be biased downwards.

Previous studies investigating the interest rate sensitivity of investment or, more generally, the sensitivity of investment to the user cost of capital, find mixed results (using methods other than those in our paper). *Guiso* et al. (2002) and *Gilchrist/Zakrajsek* (2007) report relatively high long-term elasticities. *Ca-*

[2] The user cost of capital depends not only on the interest rate, but also on the rate of depreciation and the relative price of investment to output (*Jorgenson* 1963). The size of the effect of the real user cost on investment is influenced by the elasticity of substitution between capital and labor. The higher the elasticity of substitution, the stronger the decline in investment after an increase in the user cost.

ballero (1999) and *Schaller* (2006) also find a negative effect of the user cost on investment. However, at the same time, several studies report that the interest rate may be less important than are quantity variables in influencing investment decisions (for an early review, see, e.g. *Chirinko* 1993). In addition, results from a survey among firms do not imply a large interest rate sensitivity of investment (*Sharpe/Suarez* 2013).

Recently, it has been argued that the interest rate sensitivity of the overall economy has declined (see, among others, *Willis/Cao* 2015). Provided that this argument also applies to investment, capital accumulation in advanced economies may have become less responsive to monetary policy. As a result, central banks lowering their policy rates would have become less successful in stimulating investment than in the past and, at the same time, less successful in curbing investment booms. This raises questions about the real effects of monetary policy shocks. If the interest rate under consideration is the short-term interest rate targeted by the central bank, interest rate shocks may reflect monetary policy shocks to a significant extent. There is a large body of literature investigating monetary policy shocks using VAR models (see, e.g. *Bernanke/Blinder* 1992; *Bernanke/Mihov* 1995; *Bagliano/Favero* 1998 or *Uhlig* 2005). Several papers, for instance, *Peersman* (2004) and *Boeckx/DosschePeersman* (2017), have investigated differences across countries with respect to the effects of monetary policy shocks. In contrast to most of these studies that study the effects of shocks to the interest rate (or monetary aggregates) on inflation and output, our analysis is focused on how investment responds to interest rate shocks. In VAR models, the challenge for researchers is to disentangle policy makers' responses to nonmonetary developments from monetary innovations. Endogenous changes are reactions of the interest rate to the evolution of macroeconomic variables, whereas exogenous policy captures all other actions and represent fundamental shocks to the economy. Thus, the better we are at disentangling endogenous from exogenous shifts in the interest rate, the more successful we can be at addressing the simultaneity issue between interest rates and investment.

An exogenous interest rate shock may be interpreted in a variety of ways (see, e.g. *Christiano* et al. 1999 or *Uhlig* 2005). First, one may interpret it as an exogenous shock to the preferences of central bankers as, for example, shifts in the weights given to inflation or the output gap. Second, exogenous shifts may occur because a central bank tries to meet the expectations of financial market participants. Shocks to these expectations will translate to shocks in the interest rate policy of central banks. A third reason for the detection of interest rate shocks in a VAR model is measurement errors in real-time data that are available when a central bank decides on the level of interest rates. Fourth, changes in the interest rate may reflect market forces as well as policy decisions. Central banks usually allow small movements in their targeted short-term interest rate that are driven by market forces.

The rest of this paper is organized as follows. In Section 2, we describe the data and methodology used in this paper. In Section 3, we present the results of our analysis. Finally, Section 4 contains the conclusion.

II. Data and Methodology

1. Data

We use the following variables in our VAR analyses.[3] I_t is the log of real gross fixed capital formation. GDP_t stands for the log of the real gross domestic product. P_t is the log of the price level. For the United States and the United Kingdom, we use the core price index without food and energy. For the other countries, we use the price index with the longest available time series. For Germany and Italy, this is the consumer price index, and, for France, it is the GDP deflator. Finally, ir_t is the short-term interest rate. For the United States, this is represented by the federal funds rate. For the United Kingdom, the bank rate is chosen, which is the interest rate set by the Bank of England. We use the three-month money market rate for the remaining European countries in order to have long time series. For a robustness check, we also use the long-term interest rates for government bonds with a maturity of ten years for all countries.

For the United States, data availability allows us to use time series data starting in 1957 Q1. The data for the United Kingdom start in 1965 Q1. For the other European countries, we use data starting in 1970 Q1. These data provide us with fairly long time series. Earlier data were not available for at least one time series in these countries. In the baseline regressions, we use data that end in 2007 Q4 to exclude the period since the financial crisis that could lead to misleading results, as interest rates have remained very low, and quantitative easing measures have supplemented considerably the interest rate channel of monetary policy. However, we also conduct sensitivity analysis using data that end in 2018 Q3.

2. Methodology

a) The VAR Model

As discussed above, we use a VAR model of order p of the form

$$y_t = A_1 y_{t-1} + \ldots + A_p y_{t-p} + u_t\,, \qquad \textit{with} \quad u_t \sim \mathcal{N}(0, \Sigma_u)$$

[3] Data for the United States comes from the St. Louis Fed and the Bureau of Economic Analysis. For European countries, we use data provided by Eurostat and the national central banks.

where $y_t = (y_{1t}, ..., y_{Kt})'$ is a vector of observable variables, the A_i are $(K \times K)$ coefficient matrices and u_t are K-dimensional serially uncorrelated reduced form residuals. The vector of variables in logs is given by $y_t = (I_t, GDP_t, P_t, ir_t)'$. I_t stands for real private fixed investment, GDP_t is real gross domestic product, P_t is the price level, and ir_t denotes the short-term interest rate. Using augmented Dickey-Fuller tests and tests on trend stationarity (*Dickey/Fuller* 1979), we find that all variables are $I(1)$. Therefore, we express our variables in first differences for the empirical analysis. The structural residuals ε_t in a VAR model can be obtained by pre-multiplying the matrix B, which contains the instantaneous effects of the structural shocks on the observed variables. Therefore, the structural shocks are a linear transformation of the reduced form residuals:

$$u_t = B\varepsilon_t \quad or \quad \varepsilon_t = B^{-1}u_t$$

We use a Cholesky decomposition of the covariance matrix to identify the structural shocks of our model. This choice is based on approaches followed in previous studies, which assumed that a monetary policy shock has no immediate effect on macroeconomic quantities and inflation (see, e.g. *Christiano* et al. 1999). However, this identification strategy is not uncontroversial. Since we use quarterly data, while many papers use monthly (although partly interpolated) data, the validity of our identification strategy is not guaranteed. To verify our identification, we use the identification by changes in volatility as described in *Rigobon* (2003) and *Lanne/Lütkepohl* (2008b). In this approach, we exploit the fact that we can identify a structural break for each country using Chow tests. This provides us with two distinguishable volatility regimes in the data. Importantly, this identification scheme allows us to test for overidentifying restrictions, i.e., whether our identification of the interest rate shock is supported by the data. The split samples for the pre- and post-break periods can then be used to run a VAR(p) for each period and compute impulse responses. The lag order p in our model is determined by the Akaike Information Criterion (AIC).

b) Using Chow Tests to Detect Structural Breaks

As discussed above, after estimating the VAR models for the entire period using a Cholesky decomposition, we use the changes in volatility approach to test for the appropriateness of this identification strategy and to derive separate impulse responses using this approach. This requires determining a structural break point in the parameters of the covariance matrix. To this end, we use two types of Chow tests as described in *Doornik/Hendry* (1997). The first ("break point") tries to detect a break point by testing for constant parameters and a changing covariance structure (*Hansen* 2003). The second ("sample split") tests only for constant parameters and assumes no changes in the white noise error term. These two tests help us investigate whether a structural break is due to structural changes either in the parameters or in the covariance matrix.

The first version of the Chow test (break-point "bp") tests both parameter constancy and the constancy of the white noise variance:

$$\lambda_{bp} = (T_1 + T_2)\ln\left|\hat{\Sigma}_{(1,2)}\right| - T_1 \ln\left|\hat{\Sigma}_{(1)}\right| - T_2 \ln\left|\hat{\Sigma}_{(2)}\right| \approx \chi^2,$$

where $T_1 < T_B$ and $T_2 \le T - T_B$ and $\hat{\Sigma}_{(1,2)} = \frac{1}{T_1}\sum_{t=1}^{T_1} \hat{u}_t \hat{u}_t' + \frac{1}{T_2}\sum_{t=T-T_2+1}^{T} \hat{u}_t \hat{u}_t'$, $\hat{\Sigma}_{(1)} = \frac{1}{T_1}\sum_{t=1}^{T_1} \hat{u}_t^{(1)} \hat{u}_t^{'(1)}$, $\hat{\Sigma}_{(2)} = \frac{1}{T_2}\sum_{t=T-T_2+1}^{T} \hat{u}_t^{(2)} \hat{u}_t^{'(2)}$.

The residuals $\hat{u}_t$ are obtained by running a VAR over the entire period T, whereas the residuals $\hat{u}_t^{(1)}$ and $\hat{u}_t^{(2)}$ are obtained by running a VAR over T_1 and T_2, respectively. The null hypothesis is no structural break in the parameters and the covariance structure. Since Chow test showed that the χ^2 distribution is a poor approximation, the null hypothesis is rejected too often. To overcome this problem, we use a residual based bootstrap procedure to obtain empirical quantiles.

The second Chow statistic tests for a sample split (ss) and is given by:

$$\lambda_{ss} = (T_1 + T_2)\left(\ln\left|\hat{\Sigma}_{(1,2)}\right| - \ln\left[\frac{1}{T_1 + T_2}\left(T_1 \hat{\Sigma}_{(1)} + T_2 \hat{\Sigma}_{(2)}\right)\right]\right) \approx \chi^2$$

The null hypothesis is that there is no structural break.

c) Testing the Validity of the Identifying Strategy

We attempt to test whether our identification of the interest rate shock is justified by the data. To this end, we use the changes in volatility approach, as in *Rigobon* (2003) and *Lanne/Lütkepohl* (2008*b*). Following this approach, we start by considering the two covariance matrices obtained after imposing the structural break in period T_B:

$$\mathbb{E}[u_t u_{t'}] = \begin{cases} \Sigma_1 & for \quad t = 1,\dots,T_B - 1 \\ \Sigma_2 & for \quad t = T_B,\dots,T, \end{cases}$$

The two matrices can be decomposed into $\Sigma_1 = BB'$ and $\Sigma_2 = B\Psi B'$. As usual, B is the matrix of the instantaneous shock responses and is a diagonal matrix capturing the changes in volatility across the two periods (*Lanne/Lütkepohl* 2008a). We use the break point found by the Chow test as an exogenous break point. This ensures that the break points are chosen carefully. However, even if the break points are fixed incorrectly, the time invariant parameters can still be estimated consistently (*Rigobon* 2003). The covariance matrices are given by:

$$\hat{\Sigma}_1 = \frac{1}{T_B - 1}\sum_{t=1}^{T_B - 1} \hat{u}_t \hat{u}_t^{'} \ ; \ \hat{\Sigma}_2 = \frac{1}{T - T_B + 1}\sum_{t=T_B}^{T} \hat{u}_t \hat{u}_t^{'}$$

The estimates $\hat{B}$ and $\hat{\Psi}$ can be obtained by maximizing the log likelihood:

$$\ln \mathcal{L} = -\frac{T_B - 1}{2}\left(\ln|BB'| + tr\ (\hat{\Sigma}_1 (BB')^{-1})\right) - \frac{T - T_B + 1}{2}\left(\ln|B\Psi B'| + tr\ (\hat{\Sigma}_2 (B\Psi B')^{-1})\right).$$

Moreover, we use the procedure introduced by *Lanne/Lütkepohl* (2008b) to improve the estimation precision of this method. The matrices $\hat{B}$ and $\hat{\Psi}$ obtained from maximizing the log likelihood function are used in an iterative GLS estimation. The GLS method uses $\hat{\beta}$ to update the covariance estimates by $\hat{u}_t = y_t - (Z_t^{'} \otimes I_K)\hat{\beta}$.

$$\begin{aligned}\hat{\beta} &= vec\left[\hat{\nu}, \hat{A}_1, \ldots, \hat{A}_p\right] \\ &= \left[\sum_{t=1}^{T_B - 1}\left(Z_t Z_t^{'} \otimes (\hat{B}\hat{B}')^{-1}\right) + \sum_{t=T_B}^{T}\left(Z_t Z_t^{'} \otimes (\hat{B}\hat{\Psi}\hat{B}')^{-1}\right)\right]^{-1} \\ &\quad \times \left[\sum_{t=1}^{T_B - 1}\left(Z_t \otimes (\hat{B}\hat{B}')^{-1}\right) y_t + \sum_{t=T_B}^{T}\left(Z_t \otimes (\hat{B}\hat{\Psi}\hat{B}')^{-1}\right) y_t\right],\end{aligned}$$

where $Z_t^{'} = \left[y_{t-1}', \ldots, y_{t-p}'\right]$. This procedure iterates until convergence of the likelihood. To derive standard errors for the estimates, we calculate the square root of the elements of the inverted Fisher information matrix (*Hamilton* 1994). The initial matrix B is the Choleski decomposition of $\hat{\Sigma}_u$, which is obtained from least squares estimation of the reduced form VAR and the initial Ø matrix is an identity matrix.

A necessary condition for identification with changes in volatility is that the elements of the main diagonal of $\hat{\Psi}$ are distinct. There seems to be a consensus in the literature for using a type of a Wald test for pairwise comparison *Lütkepohl/Netsunajev* (2014):

$$\lambda_W = \frac{(\hat{\psi}_i - \hat{\psi}_j)^2}{Var\left[\hat{\psi}_i\right] + Var\left[\hat{\psi}_j\right] - 2\, Cov\left[\hat{\psi}_i, \hat{\psi}_j\right]} \sim \chi_1^2 \qquad \forall\ i \neq j$$

The null hypothesis is that the elements are not distinct (see, for instance, *Herwartz/Ploedt* 2016 or *Lütkepohl/Netsunajev* 2014). If the elements are distinct, the model is identified. Thus, further restrictions that may be relevant from an economic perspective are overidentifying and can be tested. Therefore, we are in a position to test whether our previous identification of the interest rate shock is supported by the data, i. e., we can test our restricted model against the unrestricted model solely identified by changes in volatility. A comparable setup is used by *Lütkepohl/Netsunajev* (2014). To this end, a likelihood ratio test is used:

$$\lambda_{LR} = 2\left[\ln l\left(\tilde{\delta}\right) - \ln l\left(\tilde{\delta}_r\right)\right] \sim \chi_k^2 ,$$

where $\tilde{\delta}$ is the unrestricted and $\tilde{\delta}_r$ the restricted ML estimator. The χ^2 distribution has as many degrees of freedom as there are distinct linear restrictions (*Lütkepohl* 2007). The null hypothesis is that the unrestricted and the restricted models are equal.

III. Results

1. Testing the Appropriateness of the Identification Strategy

As discussed above, we want to test whether our identification of the interest rate shock is supported by the data. First, we estimate the unrestricted models with changes in volatility and check whether the data contain enough heteroscedasticity to identify the B matrix with changes in volatility. In this baseline version, we use data that end in 2007 Q4 as discussed above. The Akaike information criterion (AIC) is used to determine the lag-order of the VARs for each country. For the United States, the United Kingdom, France, Germany, and Italy, we get lag orders of 4, 5, 4, 3, and 4, respectively. Since we need one break point for this identification strategy, we first determine the break dates using Chow tests. The break dates should be chosen carefully, although the time invariant parameters can still be estimated consistently when the break points are fixed incorrectly (*Rigobon* 2003). We select a candidate break point for each country and introduce a period of uncertainty of three years to compute Chow test statistics for each potential break point within this time frame. The largest statistic obtained in this way provides a stability test for an unknown break point. The candidate break points are chosen based on likely break points related to changes in either monetary policy or the economic environment. In doing so, we must include a sufficient number of observations on both sides of the structural breaks. Based on these considerations, we choose the following potential break points for each country. For the United States, we follow the previous literature and select a candidate break point at 1984 Q4 (see also *Willis/Cao* 2015). For the European countries, we choose candidate break points at 1990 Q4. The German reunification officially occurred at the beginning of 1990 Q4 and may be seen as a potential breakpoint, because it led to economic and inflationary divergences in Europe that resulted in speculative attacks on the European exchange rate mechanism. Eventually, European countries had to adopt very wide exchange rate margins.

The break points detected by our tests are shown in Table 1. If the p-values differ between the tests, one can infer that the break tends to be driven either by changes in the parameters or by changes in the covariance matrix. For instance, the break for the UK seems to be driven mainly by the changing covariance structure, whereas the break for Italy is caused by changing parameter coefficients. For almost every country, the break point test indicates a structural break at a significance level of at least 10 %. The exception is Italy where neither the break point nor the sample split test yields a break point at a level of at least 10 %. The potential break point with the lowest p-value is at 1996 Q3. The sample split test yields a p-value of 14.5 % for this break point. Based on these break points, we further analyze whether there is sufficient heteroscedasticity in the data for identifying the structural shocks with the changes in volatility approach.

Table 1

Time Breaks in Every Country with p-Values of Chow Tests Based on 2000 Bootstrap Replications

Country	Time Break Date	λ_{bp} p-value	λ_{ss} p-value
USA	1984 Q4	< 0.001	0.616
UK	1991 Q3	0.023	0.981
France	1991 Q4	0.024	0.412
Germany	1988 Q2	0.098	0.493
Italy	1996 Q3	0.372	0.145

Based on these structural breaks, we then investigate whether the elements of the main diagonal of $\hat{\Psi}$ are distinct, i. e., we check whether the pre- and post-break periods are characterized by changes in the volatility of the shocks. The results are shown in Table 2. The null hypothesis of equal elements in $\hat{\Psi}$ cannot be rejected for all elements in this matrix, which means that the shocks are only partially identified. Since our analysis focuses on the interest rate shock, we are allowed to proceed with all countries and estimate restricted models to identify at least the interest rate shock. Our partial identification and the focus on the interest rate shock also imply that the ordering of the variables investment, GDP, and inflation is not relevant. Table 3 shows the estimated $\hat{B}$ and $\hat{\Psi}$ matrices. These matrices are obtained through restricted maximum likelihood estimation, where the shocks are identified through changes in volatility and zero restrictions. Almost every element in the $\hat{\Psi}$ matrices is smaller than 1, indicating that the countries under investigation generally shifted to a regime of lower volatility after the structural break, with the interest rate shock in the second time period characterized by a lower volatility for all countries. For instance, the fourth column of the $\hat{\Psi}$ matrix for the UK implies that the variance of a monetary policy shock in the second regime is only 13.6 % as high as is that in the first regime.

Next, the unrestricted model identified by changes in volatility (M1) is tested against the model identified with zero restrictions via the Cholesky decomposition (M2). We first test whether M2 is rejected. If it is not, we gain confidence that our identification scheme is appropriate. However, if model M2 is rejected, we proceed and test whether the identification of the structural interest rate shock cannot be rejected. To this end, we define model M3, in which zero restrictions in the B matrix are self-imposed by the optimized coefficients in B. Zero restrictions are self-imposed when the 95 % quantiles around the optimized coefficients include zero. If M3 is rejected, we will have to impose additional restrictions. However, this is unnecessary, as neither M2 nor M3 is rejected for all countries. Most of the models support the zero restriction of a Cholesky decomposition. Only the $\hat{B}$ matrices for France do not; hence only the interest rate shock is identified.

Table 2

Wald Statistic with p-Values from Pairwise Comparison of the Elements of the Unrestricted Models

Country/H_0	$\psi1 = \psi2$	$\psi1 = \psi3$	$\psi1 = \psi4$	$\psi2 = \psi3$	$\psi2 = \psi4$	$\psi3 = \psi4$
USA	0.05	< 0.001	0.02	< 0.001	< 0.001	<0.001
UK	< 0.001	< 0.001	< 0.001	< 0.001	< 0.001	0.1
France	< 0.001	0.01	0.09	0.03	< 0.001	0.22
Germany	< 0.001	0.07	< 0.001	0.07	0.09	<0.001
Italy	0.02	0.27	0.27	< 0.001	0.13	0.04

Table 3

Estimations of B and Ψ with Changes in Volatility and Zero Restrictions of the Model Which is the Closest to a Cholesky Decomposition, But Not Rejected by a Likelihood Ratio Test

Country	$\widehat{B}$	$\widehat{\Psi}$	LR test H_0	LR test p-value
USA	$\begin{bmatrix} \underset{(0.144)}{2.030} & 0 & 0 & 0 \\ \underset{(0.056)}{0.461} & \underset{(0.046)}{0.642} & 0 & 0 \\ \underset{(0.017)}{0.003} & \underset{(0.018)}{-0.051} & \underset{(0.037)}{0.522} & 0 \\ \underset{(0.830)}{2.745} & \underset{(0.831)}{0.864} & \underset{(0.944)}{2.457} & \underset{(0.673)}{9.414} \end{bmatrix}$	$\begin{bmatrix} \underset{(0.067)}{0.329} & 0 & 0 & 0 \\ 0 & \underset{(0.059)}{0.286} & 0 & 0 \\ 0 & 0 & \underset{(0.007)}{0.035} & 0 \\ 0 & 0 & 0 & \underset{(0.159)}{0.770} \end{bmatrix}$	M2	0.313
UK	$\begin{bmatrix} \underset{(0.193)}{2.659} & 0 & 0 & 0 \\ \underset{(3.684)}{4.477} & \underset{(11.871)}{163.213} & 0 & 0 \\ \underset{(0.044)}{-0.058} & \underset{(0.091)}{-0.221} & \underset{(0.067)}{0.928} & 0 \\ \underset{(0.512)}{0.865} & \underset{(1.137)}{2.333} & \underset{(0.883)}{0.227} & \underset{(0.863)}{11.872} \end{bmatrix}$	$\begin{bmatrix} \underset{(0.206)}{0.915} & 0 & 0 & 0 \\ 0 & \underset{(0.007)}{0.033} & 0 & 0 \\ 0 & 0 & \underset{(0.039)}{0.174} & 0 \\ 0 & 0 & 0 & \underset{(0.031)}{0.136} \end{bmatrix}$	M2	0.889
France	$\begin{bmatrix} * & * & * & 0 \\ * & * & * & 0 \\ * & * & * & 0 \\ * & * & * & \underset{(0.829)}{10.393} \end{bmatrix}$	$\begin{bmatrix} * & 0 & 0 & 0 \\ 0 & * & 0 & 0 \\ 0 & 0 & * & 0 \\ 0 & 0 & 0 & \underset{(0.136)}{0.57} \end{bmatrix}$	M2 M3	0.05 0.581
Germany	$\begin{bmatrix} \underset{(0.094)}{1.106} & 0 & 0 & 0 \\ \underset{(0.161)}{0.983} & \underset{(0.207)}{2.428} & 0 & 0 \\ \underset{(0.022)}{-0.007} & \underset{(0.029)}{0.001} & \underset{(0.026)}{0.303} & 0 \\ \underset{(0.615)}{1.993} & \underset{(0.919)}{1.471} & \underset{(0.719)}{1.470} & \underset{(1.094)}{12.702} \end{bmatrix}$	$\begin{bmatrix} \underset{(0.382)}{1.638} & 0 & 0 & 0 \\ 0 & \underset{(0.124)}{0.530} & 0 & 0 \\ 0 & 0 & \underset{(0.242)}{1.034} & 0 \\ 0 & 0 & 0 & \underset{(0.076)}{0.319} \end{bmatrix}$	M2	0.3
Italy	$\begin{bmatrix} \underset{(0.110)}{1.570} & 0 & 0 & 0 \\ \underset{(0.054)}{0.263} & \underset{(0.048)}{0.682} & 0 & 0 \\ \underset{(0.932)}{-0.228} & \underset{(0.977)}{0.502} & \underset{(0.753)}{10.727} & 0 \\ \underset{(0.820)}{-1.067} & \underset{(0.880)}{2.461} & \underset{(0.744)}{3.011} & \underset{(0.66)}{9.338} \end{bmatrix}$	$\begin{bmatrix} \underset{(0.204)}{0.811} & 0 & 0 & 0 \\ 0 & \underset{(0.116)}{0.459} & 0 & 0 \\ 0 & 0 & \underset{(0.286)}{1.128} & 0 \\ 0 & 0 & 0 & \underset{(0.179)}{0.694} \end{bmatrix}$	M2	0.235

Standard Deviations are in Brackets

2. *Impulse Responses for the Two Subperiods*

Figure 1 shows impulse response functions for each country for the pre- and post-break period for the sample that ends in 2007 Q4. The percentage change in the level of investment to an unexpected 25 basis point cut in the interest rates before and after the break point is depicted.[4] 68 % confidence bands are based on 2000 bootstrap replications.

[4] As discussed above, we use the first-differenced variables in logs in the VAR models.

In general, the impulse response functions show that the interest rate sensitivity of investment is lower in the second subperiod than in the first. For most countries, the responses imply that investment shows a positive reaction to a cut in the interest rate in the first subperiod ending in the 1980s or 1990s. In the second sub-period, the reaction is more dampened in all countries, and confidence bands frequently include the zero line. Overall, one may conclude based on our results that investment in the second sub-period barely responds to a cut in the interest rates. For some countries, interest rate shocks may even have a negative effect on investment according to the results of this particular model. The results reveal interesting differences between the United States and Europe, particularly for the first subperiod. In this subperiod, the interest rate sensitivity of investment was particularly strong in the United States, where the response implies an approximately 4 % increase in investment after 1.5 years. In Germany and Italy, however, the response is considerably lower, with increases after 1.5 years of less than two percent. The United Kingdom and France lie somewhere in between. Interestingly, in the second subperiod, differences between the United States and Europe mostly disappear.

In a sensitivity analysis, we show impulse response functions for the sample that ends in 2018 Q3, but with the same break date (Figure 2). The results show that the qualitative pattern stays the same as for the sample that ends before the financial crisis. For Germany, however, the difference between the first and second subperiod is significantly less pronounced. For Italy, the impulse responses show a stronger decrease in the interest rate sensitivity of investment in the second subperiod. In a further sensitivity analysis, we use the long-term interest rate for government bonds with a maturity of ten years rather than the short-term interest rate. This does not alter the qualitative pattern of our results.[5]

A number of potential factors may have led to the decline in the interest rate sensitivity of investment. First, financial market innovations and regulatory changes could have weakened the transmission mechanism of monetary policy via interest rates. Second, it can be argued that global rather than domestic interest rates increasingly determine the relevant interest rates for domestic firms. This weakens the interest rate channel of domestic monetary policy, especially in small open economies (*Hoerdahl/Sobrun/Turner* 2016; *Gudmundsson* 2017). Third, structural transformations, namely the shift from manufacturing to less interest rate sensitive services could have led to the overall decline in the interest rate sensitivity of investment.

[5] The results can be obtained from the authors upon request.

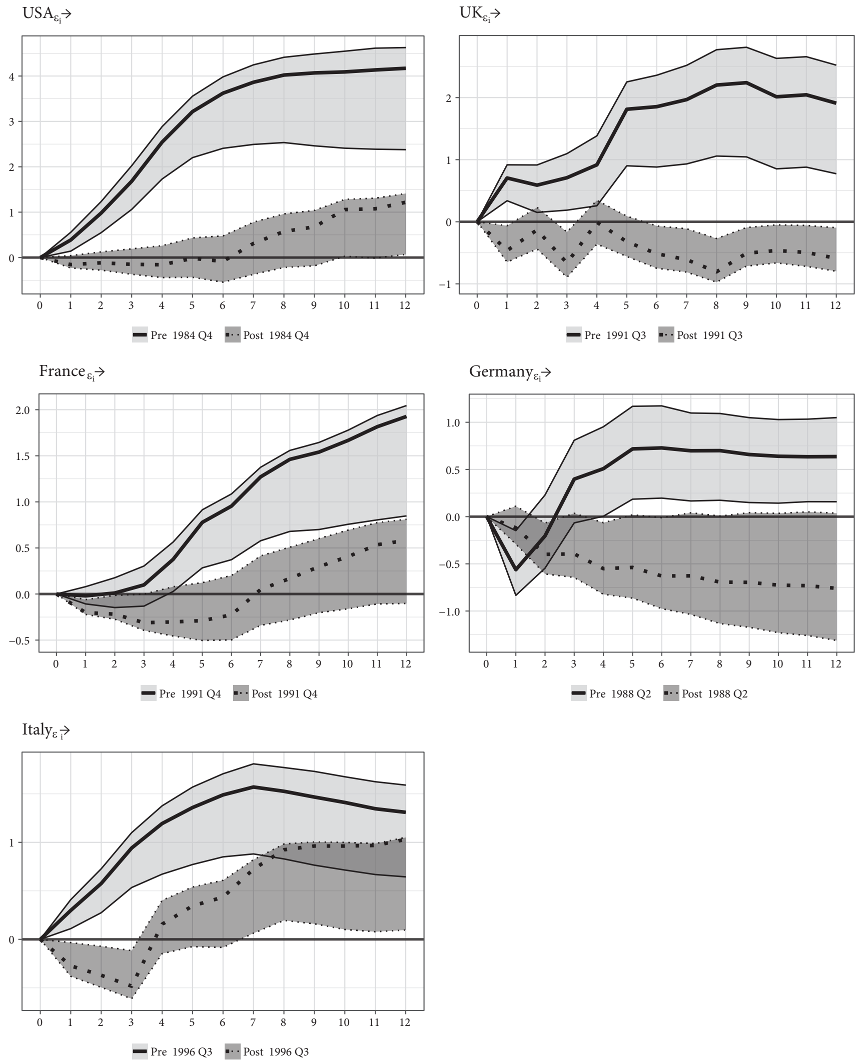

Figure 1: Impulse Response Functions of the Percentage Change in Investment to an Unexpected 25 Basis Point Cut in the Interest Rates Before and After the Break Point (Data end in 2007 Q4). 68% Confidence Bands are Based on 2000 Bootstrap Replications

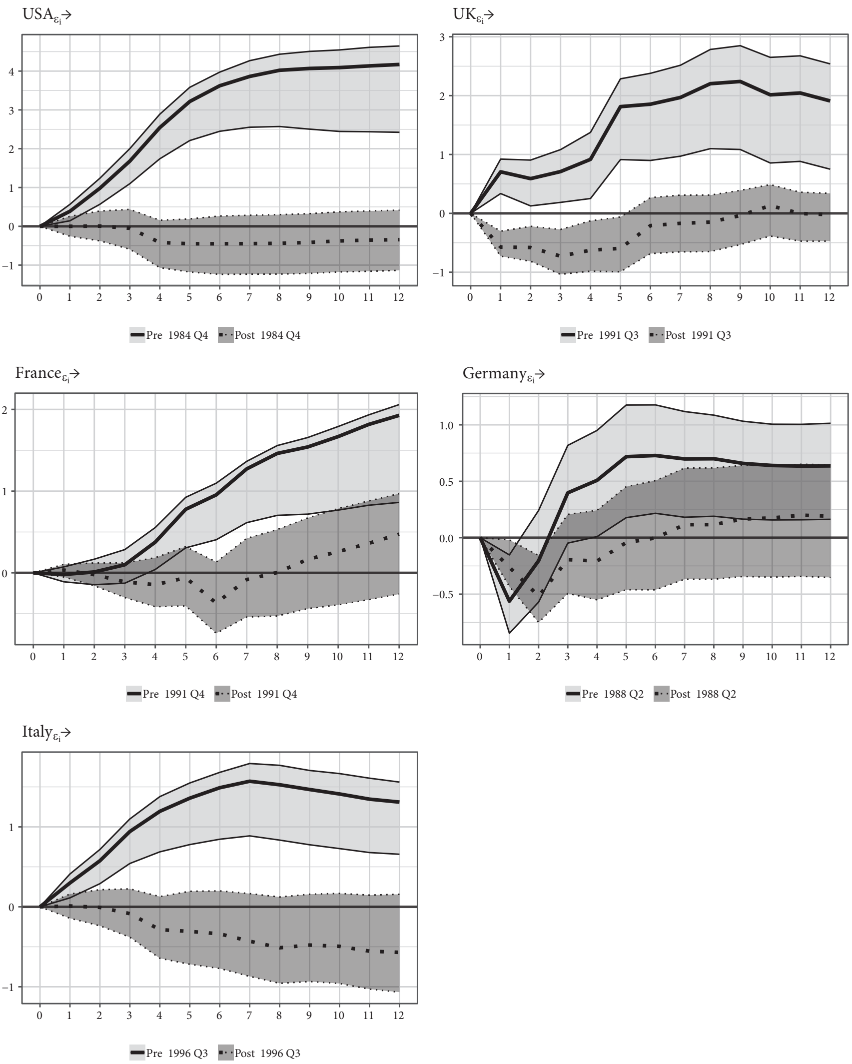

Figure 2: Impulse Response Functions of the Percentage Change in Investment to an Unexpected 25 Basis Point Cut in the Interest Rates Before and After the Break Point (Data end in 2018 Q3). 68 % Confidence Bands are Based on 2000 Bootstrap Replications

IV. Conclusion

In this paper, we empirically investigate the interest rate sensitivity of investment for the United States and the four largest European economies. In particular, we analyze whether this sensitivity has declined in recent decades. We use a vector autoregressive model with four variables: real investment, real gross domestic product, inflation, and a measure of the short-term interest rate. In our model, the structural interest rate shock is identified assuming that macroeconomic quantities and inflation react with a lag to interest rate innovations. The appropriateness of this specification is tested using the changes in volatility approach, exploiting the evidence that the volatilities of the shocks differ across two sub-periods. This approach might be useful for future empirical research in economic history, where changes in volatility may be particularly frequent. Using Chow tests to determine specific break points, we split the sample into pre- and post-break periods. For the countries under consideration, we detect a period during either the 1980s or the 1990s where the interest rate sensitivity of investment started to decrease. According to our findings, expansionary interest rate shocks have had expansionary effects on real investment in the decades before the 1980s, this seems to be less the case more recently. Our results suggest that, in recent decades, interest rate shocks have displayed ambiguous real effects on investment and may be neutral. The decrease in the interest rate sensitivity of investment is particularly pronounced for the United States. The past three decades have been characterized by a decline in the level of interest rates and a subdued evolution of corporate investment. The findings in this paper suggest that these long-term developments are not influenced by monetary policy, but are driven by structural factors. Future research could investigate the robustness of our results using different data and empirical methods.

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