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Three Essays on Empirical Analysis of Economic Policy

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I hereby declare that this dissertation thesis is my original work and that all sources and literature used are listed. This dissertation thesis has not been used to acquire a different or the same university degree.

Signature

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Abstract

This dissertation thesis is focused on the empirical analysis of monetary and fiscal policy using nonlinear models. It consists of three parts, the first two parts deal with the analysis of monetary policy using the monetary policy rule with time-varying parameters. The third part of this thesis is focused on finding answer to the question, whether the negative effects of financial instability on economic growth can be mitigated by expansionary fiscal policy.

In the first part, I examine the evolution of monetary policy rules in a group of inflation targeting countries (Australia, Canada, New Zealand, Sweden and the United Kingdom). I apply a moment-based estimator in a time-varying parameter model with endogenous regressors. Using this novel flexible framework, the main findings are threefold. First, with adoption of inflation targeting, coefficients in the monetary policy rules changed rather gradually, pointing to the importance of applying a time-varying estimation framework. Second, the interest rate smoothing parameter is much lower than typically reported by previous time-invariant estimates of policy rules. Third, the response of interest rates to inflation is particularly strong during periods when central bankers want to break a record of high inflation, such as in the UK or Australia at the beginning of the 1980s. Contrary to common perceptions, the response of interest rates to inflation becomes less aggressive after the adoption of inflation targeting, suggesting a positive anchoring effect of this regime on inflation expectations. This result is supported by our finding that inflation persistence typically decreased after the adoption of inflation targeting.

The second part discusses whether and how the selected central banks responded to episodes of financial stress over the last three decades. The time-varying monetary policy rule is extended for an indicator of financial stress, in order to show the departures of policy rules under financial instability. To measure the financial stress, I use a new financial stress dataset developed by the International Monetary Fund. This particular choice not only allows testing of whether central banks responded to financial stress or not, but also detects the periods and types of stress that were the most worrying for monetary authorities and quantifies the intensity of the policy response. The findings suggest that central banks often change policy rates, mainly decreasing them, in the face of high financial stress. However, the size of the policy response varies substantially over time as well as across countries, with the 2008–2009 financial crisis being the period of the most severe and generalized response. With regard to the specific components of financial stress, most central banks seemed to respond to stock-market stress and bank stress, while exchange-rate stress is found to drive the reaction of central banks only in more open economies.

In the third part, I use a threshold VAR model to study whether the effects of fiscal policy on economic activity differ depending on financial market conditions. In particular, I investigate the possibility of a non-linear propagation of fiscal developments according to different financial market stress regimes. More specifically I employ a quarterly dataset, for the U.S., the U.K., Germany and Italy, for the period 1980:4-2009:4, encompassing macro, fiscal and financial variables. The results show that output reacts mostly positively to a fiscal shock in both financial stress regimes, and differences in estimated multipliers across regimes are relatively small. The large time-variation and the estimated nonlinear impulse responses suggest that the size of the fiscal multipliers is higher than average in the 2008-2009 crisis. Furthermore, a financial stress shock has a negative effect on output and worsens the fiscal situation.

Abstrakt

Tato disertační práce je zaměřena na empirickou analýzu měnové a fiskální politiky s využitím nelineárních modelů. Skládá se ze tří částí, první dvě části se věnují analýze měnové politiky pomocí odhadnuté reakční funkce centrálních bank s časově proměnlivými parametry. Třetí část předložené práce je zaměřena na hledání odpovědí na otázku, jestli je možné finanční nestabilitu a související nízký růst HDP překonat pomocí expanzivní fiskální politiky.

V první části se zabývám analýzou vývoje měnové politiky v zemích s inflačním cílováním (Austrálie, Kanada, Nový Zéland, Švédsko a Británie). Díky využití nelineárního přístupu, tj. modelu s parametry, které se v čase mění, můžeme zkoumat, jak velký dopad mělo zavádění inflačního cílování na skutečnou politiku centrálních bank a částečně můžeme i ukázat, jak inflační cílování ovlivnilo dynamiku inflace. Z pohledu metodologie analýza stojí na empirickém odhadu měnového pravidla s endogenními regresory a s koeficienty, jejichž dynamika sleduje náhodnou procházku. Samotný odhad je proveden pomocí momentového odhadu. Ukazuji, že zavedení inflačního cílování není v odhadnutých parametrech patrné jako náhlý zlom, ale jejich dynamika spíše ukazuje, že zavedení inflačního cílování byl postupný proces. V tomto ohledu se projevuje přednost zvoleného přístupu, které umožňuje zachytit jak náhlé strukturální změny, tak postupné změny dynamiky. Za druhé, mé výsledky vedou k nízkým hodnotám vyhlazování úrokové míry, zejména ve srovnání s modely s konstantními koeficienty. A konečně, výsledky neukazují, že by inflační cílování bylo provázeno restriktivnější měnovou politikou, ale spíše naopak. Stabilní inflace je tak spíše důsledkem vyšší kredibility centrálních bank a schopností ovlivnit inflační očekávání, než restriktivní politikou.

Cílem druhé části je analýza, zda a jak v posledních třech desetiletích reagovaly vybrané centrální banky v obdobích finanční nestability na finančních trzích. Analýza využívá modelu odhadu měnově-politických pravidel s časově proměnlivými parametry a tato metodologie je společně s indexem finančního stresu, který byl nově vyvinut Mezinárodním měnovým fondem, aplikována na Austrálii, Kanadu, Švédsko, Velkou Británii a USA. Empirická analýza umožňuje nejen otestovat, zda centrální banky těchto zemí na finanční nestabilitu vůbec reagovaly, ale i kvantifikovat intenzitu jejich reakce. Z výsledků studie vyplývá, že období finanční nestability centrální banky často měnily měnově-politické sazby, a to převážně směrem dolů. Intenzita reakce měnové politiky je nicméně v čase i mezi zeměmi značně heterogenní, přičemž finanční krize 2008-2009 je obdobím nejsilnější a nejvíce rozšířené reakce.

Třetí část je zaměřena na fiskální politiku v době finanční nestability a nižšího nebo dokonce záporného hospodářského růstu. S pomocí nelineárního vícerovnicového modelu (tzv. Threshold VAR model) hledám odpověď na otázku, jestli je možné snížit finanční nestabilitu a zvýšit růst pomocí expanzivní fiskální politiky. Model obsahuje dva režimy. První režim je charakterizovaný stabilními finančními trhy a růstem HDP, druhý režim zahrnuje období finanční nestability. Druhý režim je identifikovaný tak, že obsahuje i všechna období spojená s recesí. Ukazuji, že model se dvěma režimy je statisticky významně odlišný od jednoduchého dynamického modelu bez nelinearity. Dále zjišťuji, že pro USA, Británii, Německo a Itálii, hospodářský růst reaguje rozdílně na fiskální politiku a že multiplikátor fiskální politiky je během současné krize významnější, než v minulých obdobích.

Contents

Chapter 1

Introduction

This dissertation thesis is focused on the empirical analysis of monetary and fiscal policies. It contains three essays, each of them is independent on the others. However, they all share an utilization of nonlinear empirical methods that capture time-varying and possibly asymmetric behaviour of macroeconomic policies and their effects.

The first two essays deal with the analysis of monetary policy under inflation targeting. The first essay discuss the effects of adoption and implementation of inflation targeting from the perspective of time variation in coefficients of the monetary policy rule. More specifically, it estimates the monetary policy rules in Australia, Canada, New Zealand, Sweden and the United Kingdom and it examines to what extent these rules change over time. To capture both potential endogeneity problem and time-variance in parameters of the monetary policy rule, a two-step modification of the time-varying parameter model is used.

The main findings can be summarized as follows. First, the results indicate that in these countries the changes in monetary policy are rather gradual and coincide either with important institutional reforms such as the changes in monetary policy regime or with the periods of disinflation. Apart from that, the results point to an importance of appropriate estimation framework for analysis of interest rate setting (the monetary policy rule) over time. It is shown that the degree of interest rate smoothing is much lower than many previous empirical contributions suggested, and the reason for overestimating the degree of interest rate smoothing can be in ignoring the time-varying nature of policy rules.

The first essay is a joint work with Roman Horváth and Bořek Vašíček and it has been published as the Czech National Bank Working Paper 2/2010, as the IES FSV UK Working Paper 26/2010, and currently, it is forthcoming in Macroeconomic Dynamics. Beside the initial discussions on intended focus of the paper, the contribution of mine is especially related to methodology (selection of method and computation) and construction of dataset. Additionally, This version contains number of additional materials, namely plots of all data series used (Appendix 2) and sections 2.4, 2.5 and Appendix 1 were partially revised and updated.

In the second essay, the analysis of time-variance in monetary policy rules is extended for the analysis of central bank policies under financial instability. It is analysed, whether the central banks under inflation targeting adjust their policy rates not only with respect to expected deviation between inflation and inflation target, but also with respect to financial stability. For this purpose, the time-varying monetary policy rule is extended for an indicator of financial stress, in order to show the departures of policy rules under financial instability. To measure the financial stress, a new financial stress index is used, developed by the International Monetary Fund. This particular choice not only allows testing whether central banks responded to financial stress or not, but also detects the periods and types of stress that were the most worrying for monetary authorities and quantifies the intensity of the policy response.

Although theoretical studies disagree about the viability of considering financial instability for interest-rate setting, the results show that monetary policy is likely to react to financial instability by decreasing their policy rates. However, the size of the policy response varies substantially over time as well as across countries, with the 2008–2009 financial crisis being the period of the most severe and generalized response. The results also point to the usefulness of augmenting the standard version of the monetary policy rule by some measure of the financial conditions to get a better understanding of the interest-rate setting process, especially when financial markets are not stable.

As well as the first essay, the second one is a joint work with Roman Horváth and Bořek Vašíček. It is currently forthcoming in Journal of Financial Stability and its earlier version has been published under the title "How Does Financial Instability Matter for Monetary Policy?" in Eijffinger, S., and Masciandaro, D. (eds.): Handbook of Central Banking, Financial Regulation and Supervision, Edward Elgar Publishing, pp. 269-313. The paper has been published as the Czech National Bank Working Paper 3/2011, too. My contribution is related namely to computations, interpretation of the results with narrative evidence on the stress periods and reactions of the monetary policy and also the idea to focus on time variation in the effect of financial stress on interest rate setting rather than the time variation of the coefficients itself.

Finally, the third essay is focused on finding answer to the question, whether the negative effects of the financial instability on economic growth can be overcome by expansionary fiscal policy. During periods of economic downturn or stress in financial markets the effects of fiscal developments on economic activity might be different from what is usually observed in good or normal times. The quality of financial institutions' assets deteriorates, as the share of non-performing loans increases and negative sentiments in the markets depress the value of other financial assets. In some cases, the disruptions in financial markets or problems in the banks' balance sheets may trigger a recession by reducing the flow of credit to the other sectors. Therefore, it is important to assess the effects of fiscal developments and policies during the periods of market stress to check, whether there are some non-linearities at play and if the fiscal multipliers are different.

The main contribution of this paper is that the effects of fiscal policy shocks are estimated using a threshold VAR model (TVAR) with two regimes, determined by a measure representing financial instability, the Financial Stress Index. According to my knowledge, there have been no attempts to investigate empirically the effects of fiscal developments associated with periods of financial crises within a multi-equation framework, which is the issue addressed in this paper.

Several results of the analysis are worthwhile mentioning. First, the use of a nonlinear framework with regime switches, determined by a financial stress indicator, is corroborated by nonlinearity tests. Second, output reacts mostly positively to a fiscal shock in both financial stress regimes, and the differences in estimated multipliers across regimes are relatively small. However, it is shown that substantial time-variation and the estimated nonlinear impulse responses suggest that the size of the fiscal multipliers is higher than average in the 2008-2009 crisis. Furthermore, a financial stress shock has a negative effect on output and increases the debt.

This essay has been carried out during my internship at the European Central Bank in cooperation with António Afonso and Michal Slavík who contributed namely with revisions and comments to the text and with section 4.4.3. It has been published as the European Central Bank Working Paper No. 1319 (April 2011) and the IES FSV UK Working Paper 16/2011.

Chapter 2

How Does Monetary Policy Change?

Evidence on Inflation Targeting Countries

2.1 Introduction

The Taylor-type regressions have been applied extensively in order to describe monetary policy setting for many countries. The research on U.S. monetary policy usually assumes that monetary policy was subject to structural breaks when the FED chairman changed. Clarida et al. (2000) claims that the U.S. inflation during the 1970s was unleashed because the FED's interest rate response to the inflation upsurge was too weak, while the increase of such response in the 1980s was behind the inflation moderation. Although there is ongoing discussion on the sources of this Great Moderation (Benati and Surico, 2009), the fact that monetary policy setting evolves over time is generally accepted.

The evolution of monetary policy setting as well as the exogenous changes in economic system over time raises several issues for empirical analysis. In particular, the coefficients of monetary policy rules estimated over longer periods are structurally unstable. The common solution used in the literature is typically a sub-sample analysis (Clarida et al. 1998, 2000). Such an approach is based on rather strong assumption that the timing of structural breaks is known, but also that the policy setting does not evolve within each sub-period. Consequently, this gives impetus to applying the empirical framework that allows for regime changes or in other words the time variance in the model parameters (Cogley and Sargent, 2001, 2005). Countries that implemented the inflation targeting (IT) regime are especially suitable for such analysis because it is likely that the monetary policy stance with respect to inflation and other macroeconomic variables changed as a consequence of the IT implementation. Moreover, there is ongoing debate to what extent the IT represents a rule-based policy. Bernanke et al. (1999) claim the IT is a framework or a constrained discretion rather than being a mechanical rule. Consequently, the monetary policy rule of IT central bank is likely to be time varying.

Our study aims to investigate the evolution of monetary policy for the countries that have had a long experience with the IT regime. In particular, we analyse the time-varying monetary policy rules for Australia, Canada, New Zealand, Sweden and the United Kingdom. As we are interested in the monetary policy evolution over relatively longer period, we do not consider countries where the IT was in place for relatively short time (Finland, Spain), or was introduced relatively recently (such as Armenia, the Czech Republic, Hungary, Korea, Norway or South Africa). We apply the recently developed time-varying parameter model with endogenous regressors (Kim and Nelson, 2006), as this technique allows us to evaluate the changes in policy rules over time, unlike Markov-switching methods does not impose sudden policy switches between different regimes. On the top of that, it also deals with endogeneity of policy rules. Unlike Kim and Nelson (2006) we do not rely on the Kalman filter that is conventionally employed to estimate time-varying models, but employ the moment-based estimator proposed by Schlicht and Ludsteck (2006)[1] for its mathematical and descriptive transparency and minimal requirements as regards initial conditions. In addition, Kim and Nelson (2006) apply their estimator to evaluate changes in U.S. monetary policy, while we focus on inflation targeting economies.

Anticipating our results, we find that monetary policy changes gradually, pointing to the importance of applying a time-varying estimation framework (see also Koop et al., 2009, on evidence that monetary policy changes gradually rather than abruptly). When the issue of endogeneity in time-varying monetary policy rules is neglected, the parameters are estimated inconsistently, even though the resulting errors are economically not large. Second, the interest rate smoothing parameter is much lower than typically reported by previous time-invariant estimates of policy rules. This is in line with a recent critique by Rudebusch (2006), who emphasizes that the degree of smoothing is rather low. External factors matter for understanding the interest rate setting process for all countries, although the importance of the exchange rate diminishes after the adoption of inflation targeting. Third, the response of interest rates to inflation is particularly strong during periods when central bankers want to break a record of high inflation, such as in the UK at the beginning of the 1980s. Contrary to common perceptions, the response can become less aggressive after the adoption of inflation targeting, suggesting a positive anchoring effect of this regime on inflation expectations or a low inflation environment. This result is consistent with Kuttner and Posen (1999) and Sekine and Teranishi (2008), who show that inflation targeting can be associated with a smaller response of the interest rate to inflation developments if the previous inflation record was favourable.

The paper is organized as follows. Section 2 discusses the related literature. Section 3 describes our data and empirical methodology. Section 4 presents the results. Section 5 concludes. An appendix with a detailed description of the methodology and additional results follows.

2.2 Related Literature

2.2.1 Monetary policy rules and inflation targeting

Although the theoretical literature on optimal monetary policy usually distinguishes between instrument rules (the Taylor rule) and targeting rules (the inflation-targeting based rule), the forward-looking specification of the Taylor rule, sometimes augmented with other variables, has commonly been used for the analysis of decision making of IT central banks. The existing studies feature great diversity of empirical frameworks, which makes the comparison of their results sometimes complicated. In the following we provide a selective survey of empirical studies aimed at the countries that we focus on.

The United Kingdom adopted IT in 1992 (currently a 2% target and a ±1% tolerance band) and the policy of the Bank of England (BoE) is subject to the most extensive empirical research. Clarida et al. (1998) analysed the monetary policy setting of the BoE in the pre-IT period, concluding that it was consistent with the Taylor rule, yet additionally constrained by foreign (German) interest rate setting. Adam et al. (2005) find by means of sub-sample analysis that the introduction of IT did not represent a major change in monetary policy conduct, unlike the granting of instrument

1 The description of this estimator is also available in Schlicht (2005), but we refer to more recent working paper version, where this estimator is described in a great detail. Several important parts of this framework were introduced already in Schlicht (1981).

independence in 1997. Davradakis and Taylor (2006) point to significant asymmetry of British monetary policy during the IT period; in particular the BoE was concerned with inflation only when it significantly exceeded its target. Assenmacher-Wesche (2006) concludes by means of a Markov-switching model that no attention was paid to inflation until IT was adopted. Conversely, Kishor (2008) finds that the response to inflation had already increased, especially after Margaret Thatcher became prime minister (in 1979). Finally, Trecroci and Vassalli (2009) use a model with time-varying coefficients and conclude that policy had been getting gradually more inflation averse since the early 1980's.

New Zealand was the first country to adopt IT (in 1990). A particular feature besides the announcement of the inflation target (currently a band of 1–3%) is that the governor of the Reserve Bank (RBNZ) has an explicit agreement with the government. Huang et al. (2001) study the monetary policy rule over the first decade of IT. He finds that the policy of the RBNZ was clearly aimed at the inflation target and did not respond to output fluctuations explicitly. The response to inflation was symmetric and a backward-looking rule does as good a job as a forward-looking one at tracking the interest rate dynamics. Plantier and Scrimgeour (2002) allow for the possibility that the neutral real interest rate (implicitly assumed in the Taylor rule to be constant) changes in time. In this framework they find that the response to inflation increased after IT was implemented and the policy neutral interest rate tailed away. Ftiti (2008) additionally confirms that the RBNZ did not explicitly respond to exchange rate fluctuations and Karedekikli and Lees (2007) disregard asymmetries in the RBNZ policy rule.

The Reserve Bank of Australia (RBA) turned to IT in 1993 (with a target of 2–3%) after decades of exchange rate pegs (till 1984) and consecutive monetary targeting.[2] De Brouwer and Gilbert (2005) using sub-sample analysis confirm that the RBA's consideration of inflation was very low in the pre-IT period and a concern for output stabilization was clearly predominant. The response to inflation (both actual and expected) increased substantially after IT adoption but the RBA seemed to consider exchange rate and foreign interest rate developments as well. Leu and Sheen (2006) find a lot of discretionality in the RBA's policy (a low fit of the time-invariant rule) in the pre-IT period, a consistent response to inflation during IT, and signs of asymmetry in both periods. Karedekikli and Lees (2007) document that the policy asymmetry is related to the RBA's distaste for negative output gaps.

The Bank of Canada (BoC) introduced IT in 1991 in the form of a series of targets for reducing inflation to the midpoint of the range of 1–3% by the end of 1995 (since then the target has remained unchanged). Demers and Rodríguez (2002) find that the implementation of this framework was distinguished by a higher inflation response, but the increase in the response to real economic activity was even more significant. Shih and Giles (2009) model the duration analysis of BoC interest rate changes with respect to different macroeconomic variables. They find that annual core inflation and the monthly growth rate of real GDP drive the changes of the policy rate, while the unemployment rate and the exchange rate do not. On the contrary, Dong (2008) confirms that the BoC considers real exchange rate movements.

Sweden adopted IT in 1993 (a 2% target with a tolerance band of 1 percentage point) just after

2 In Australia, the adoption of inflation targeting was a gradual process. As from January 1990, the RBA increased the frequency of its communications via speeches and the style of the Bank's Bulletin started to correspond to the inflation reports as introduced in New Zealand. The exact inflation target was defined explicitly later, in April 1993. Greenville (1997) describes the policy changes in Australia in great detail.

the krona had been allowed to float. The independence of Sveriges Riksbank (SR) was legally increased in 1999. Jansson and Vredin (2003) studied its policy rule, concluding that the inflation forecast (published by the Riksbank) is the only relevant variable driving interest rate changes. Kuttner (2004) additionally finds a role for the output gap, but in terms of its growth forecast (rather than its observed value). Berg et al. (2004) provide a rigorous analysis of the sources of deviations between the SR policy rate and the targets implied by diverse empirical rules. They claim that higher inflation forecasts at the early stages of the IT regime (due to a lack of credibility) generate a higher implied target from the forward-looking rule and therefore induce spurious indications of policy shocks. Their qualitative analysis of SR documents clarifies the rationale behind actual policy shocks, such as more gradualism (stronger inertia) in periods of macroeconomic uncertainty.

Finally, there are a few multi-country studies. Meirelles Aurelio (2005) analyzes the time-invariant rules of the same countries as us, finding significant dependence of the results on real-time versus historical measures of variables. Lubik and Schorfheide (2007) estimate by Bayesian methods an open economy structural model of four IT countries (AUS, CAN, NZ, UK) with the aim of seeing whether IT central banks respond to exchange rate movements. They confirm this claim for the BoE and BoC. Dong (2008) enriches their setting by incorporating some more realistic assumptions (exchange rate endogeneity, incomplete exchange-rate pass-through), finding additionally a response to the exchange rate for the RBA.

2.2.2 Time variance in monetary policy rules

The original empirical research on monetary policy rules used a linear specification with time-invariant coefficients. Instrument variable estimators such as the GMM gained popularity in this context, because they are able to deal with the issue of endogeneity that arises in the forward-looking specification (Clarida et al., 1998).[3] While a time-invariant policy rule may be a reasonable approximation when the analyzed period is short, structural stability usually fails over longer periods.

The simplest empirical strategy for taking time variance into account is to use sub-sample analysis (Taylor, 1999; Clarida et al., 2000). The drawback of this approach is its rather subjective assumptions about points of structural change and structural stability within each sub-period. An alternative is to apply an econometric model that allows time variance for the coefficients. There are various methods dealing with time variance in the context of estimated monetary policy rules.

The most common option is the Markov-switching VAR method, originally used for business cycle analysis. Valente (2003) employs such a model with switches in the constant term representing the evolution of the inflation target (the inflation target together with the real equilibrium interest rate makes the constant term in a simple Taylor rule). Assenmacher-Wesche (2006) uses the Markov-switching model with shifts both in the coefficients and in the residual variances. Such separation between the evolution of policy preferences (coefficients) and exogenous changes in the economic system (residuals) is important for the continuing discussion on the sources of the Great Moderation (Benati and Surico, 2008; Canova and Gambetti, 2008). Sims and Zha (2006) present a multivariate model with discrete breaks in both coefficients and

3 One exception is when a researcher uses real-time central bank forecasts for Taylor-type rule estimation, i.e. the data available to the central bank before the monetary policy meeting. In such case, the endogeneity problem will not arise and least squares estimation may perform well (Orphanides, 2001). However, as we will discuss in more detail below, the use of real-time data may not solve the issue of endogeneity completely.

disturbances. Unlike Assenmacher-Wesche they find that the variance of the shock rather than the time variance of the monetary policy rule coefficient has shaped macroeconomic developments in the U.S. in the last four decades.

The application of Markov-switching VAR techniques turns out to be complicated for IT countries, where the policy rules are usually characterized as forward-looking and some regressors become endogenous. The endogeneity bias can be avoided by means of a backward-looking specification (lagged explanatory variables), but this is very probably inappropriate for IT central banks, which are arguably forward-looking.[4] However, there is another distinct feature of the Markov-switching model that makes its use for the analysis of time variance in the monetary policy rule rather questionable. The model assumes sudden switches from one policy regime to another rather than a gradual evolution of monetary policy. Although at first sight one may consider the introduction of IT to be an abrupt change, there are some reasons to believe that a smooth monetary policy transition is a more appropriate description for IT countries (Koop et al., 2009). Firstly, the IT regime is typically based on predictability and transparency, which does not seem to be consistent with sudden switches. Secondly, it is likely that inflation played a role in interest rate setting even before the IT regime was introduced, because in many countries a major decrease of inflation rates occurred before IT was implemented. Thirdly, the coefficients of different variables (such as inflation, the output gap or the exchange rate) in the monetary policy rule may evolve independently rather than moving from one regime to another at the same time (see also Darvas, 2009). For instance, a central bank may assign more weight to the observed or expected inflation rate when it implements IT, but that does not mean that it immediately disregards information on real economic activity or foreign interest rates. Finally, there is relevant evidence, though mostly for the U.S., that monetary policy evolves rather smoothly over time (Boivin, 2006; Canova and Gambetti, 2008; Koop et al., 2009). Therefore, based on this research, a smooth transition seems to be a more appropriate description of reality. In a similar manner, it is possible to estimate the policy rule using STAR-type models. Nevertheless, it should be noted that STAR-type models assume a specific type of smooth transition between regimes, which can be more restrictive than the flexible random walk specification that we employ in this paper. Therefore, we leave the empirical examination of Markov-switching as well as STAR-type models for further research.[5]

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4 Psaradakis et al. (2006) proposed a solution to the endogeneity problem in the context of the Markov-switching model in the case of the term structure of interest rates.

5 We run a number of experiments on simulated data with the true coefficients containing large sudden shifts to see whether our econometric framework (see details below) delivers estimated coefficients that are gradually changing or shifting. The specification of the basic experiment was as follows: The intercept follows a slowly moving random walk (variance of innovations set to 0.3). For the independent variable we used expected inflation in the UK, with the beta coefficient set to 0.75 up to the 60th observation and 1.75 afterwards. Then we included a lag of the dependent variable with a constant coefficient equal to 0.5 and residuals with distribution N(0.15). The dependent variable was generated as the sum of these components. This example can be linked to a reaction function of a hypothetical central bank that smoothes the interest rate, that does not have the output gap in its reaction function, and that changed its aggressiveness abruptly in the middle of the sample. Then we estimated the model using the VC method and stored the value of the estimated change of the beta coefficient at the time of the switch in the data generating process. We repeated this small experiment 30 times for different sets of intercept and residuals and the estimated value of the switch was 0.98 on average, ranging from 0.9 to 1.02. The average sizes of the innovations in beta were below 0.09 in the remaining part of the samples. Clearly, in this simple setting an abrupt change in policy is detected by the model with respect to the size and timing of that change. Further experiments contained more variables in the reaction function and more switches going upwards and downwards as well. We found that, generally, sudden changes (larger than the average changes in the other time-varying coefficients) in the true coefficients resulted in switches in the estimated time-varying coefficients, too, and the varying coefficients did not

Besides simple recursive regression (e.g. Domenech et al., 2002), the Kalman filter has been employed in a few studies to estimate a coefficient vector that varies over time. Such a time-varying model is also suitable for reflection of possible asymmetry of the monetary policy rule (Dolado et al., 2004).[6] An example of such asymmetry is that the policy maker responds more strongly to the inflation rate when it is high than when it is low. Boivin (2006) uses such a time-varying model estimated via the Kalman filter for the U.S., Elkhoury (2006) does the same for Switzerland, and Trecroci and Vassalli (2010) do so for the U.S., the UK, Germany, France and Italy. However, none of these studies provides a specific econometric treatment to the endogeneity that arises in forward-looking specifications.

In this respect, Kim (2006) proposes a two-step procedure for dealing with the endogeneity problem in the time-varying parameter model. Kim and Nelson (2006) find with this methodology that U.S. monetary policy has evolved in a different manner than suggested by previous research. In particular, the Fed's interest in stabilizing real economic activity has significantly increased since the early 1990s.[7] Kishor (2008) applies the same technique for analysis of the time-varying monetary policy rules of Japan, Germany, the UK, France and Italy. He detects a time-varying response not only with respect to the inflation rate and the output gap, but also with respect to the foreign interest rate. The relevance of endogeneity correction can be demonstrated by the difference between Kishor's results and those of Trecroci and Vassalli (2010), who both study the same sample of countries.[8] The time-varying parameter model with specific treatment of endogeneity can be relevant even when real-time data are used instead of ex-post data (Orphanides, 2001). When the real-time forecast is not derived under the assumption that nominal interest rates will remain constant within the forecast horizon, endogeneity may still be present in the model (see Boivin, 2006). Moreover, this estimation procedure is also viable for reflecting measurement error and heteroscedasticity in the model (Kim et al., 2006). However, the Kalman filter applied to a state-space model may suffer one important drawback in small samples: it is rather sensitive to the initial values of parameters which are unknown. The moment-based estimator proposed by Schlicht (1981), Schlicht (2005) and Schlicht and Ludsteck (2006), which is employed in our paper and described below, allows this problem to be avoided. Moreover, it is flexible enough to incorporate the endogeneity correction proposed by Kim (2006).

2.3 Empirical Methodology

2.3.1 The empirical model

In line with Taylor (1993) most empirical studies' models assume that the central bank targets the nominal interest rate in line with the state of the economy (see Clarida et al., 1998, 2000). Such policy rule, which in the case of an IT central bank is arguably forward-looking, can be written as follows:

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imply gradual changes in these cases. Nevertheless, the ability of the time-varying parameter models to identify sudden structural breaks in parameters remains to be confirmed by careful Monte Carlo simulations, as pointed out by Sekine and Teranishi (2008).

6 Granger (2008) shows that any non-linear model can be approximated by a time-varying parameter linear model.

7 Kim et al. (2006) confirmed this finding with real-time data and additionally detected a significant decrease in the response to expected inflation during the 1990s.

8 Horváth (2009) employs the time-varying model with endogenous regressors for estimation of the neutral interest rate for the Czech Republic and confirms the importance of endogeneity bias correction terms.

$$r_t^* = \bar{r} + \beta\left(E[\pi_{t+i}|\Omega_t] - \pi_{t+i}^*\right) + \gamma E[y_{t+j}|\Omega_t] \tag{1}$$

where r_t^* denotes the targeted interest rate, $\bar{r}$ is the policy neutral rate[9], π_{t+i} stands for the central bank forecast of the yearly inflation rate i periods ahead, and π_{t+i}^* is the central bank's inflation target.[10] y_{t+j} represents a measure of the output gap. $E[\ ]$ is the expectation operator and Ω_t is the information set available at the time t when interest rates are set. Eq. (1) links the policy instrument (nominal interest rates) to a constant term (the neutral rate that would prevail if expected inflation and output were at their targeted levels), the deviation of expected inflation from its target value and the output gap.

Nevertheless, Eq. (1) is often argued to be too restrictive to provide a reasonable description of actual interest rate setting. First, it does not account for interest rate smoothing by central banks. In line with Clarida et al. (1998) most studies assume that the central bank adjusts the interest rate sluggishly to the targeted value. This can be tracked by a simple partial-adjustment mechanism:

$$r_t = \rho r_{t-1} + (1-\rho) r_t^* \tag{2}$$

where $\rho \in [0,1]$ is the smoothing parameter. Although this is line with the common wisdom that central banks are averse to abrupt changes, most studies that estimate time-invariant models find unusually high policy inertia. For instance, using quarterly data ϱ typically exceeds 0.8. Rudebusch (2006) points to an inconsistency between this finding and the practical impossibility to predict interest rate changes over a few quarters. Therefore, it is possible that the lagged dependent value takes over the impact of either autocorrelated shocks or omitted variables. The intensity of interest rate smoothing is logically reinforced in a linear time-invariant specification, as the response to some variables can be asymmetric and/or vary in time. Second, Eq. (1) assumes that the central bank aims only at the inflation rate and the output gap. However, many central banks that have implemented IT are in small open economies and may consider additional variables, in particular the exchange rate and the foreign interest rate, hence Eq. (1) can be extended for δx_t, where the coefficient δ measures the impact of additional variable x_t on interest rate setting. Therefore, in our empirical model we substitute Eq. (2) into Eq. (1), eliminating unobserved forecast variables, defining $\alpha = \bar{r} - \beta\pi_t^*$ and including additional variables, which results in Eq. (3):

$$r_t = (1-\rho)[\alpha + \beta(\pi_{t+i}) + \gamma y_t + \delta x_t] + \rho r_{t-1} + \varepsilon_t \tag{3}$$

Following Clarida et al. (1998), the intercept α can be restated in terms of equilibrium real interest rate rr such that $\alpha = rr - (1-\beta)\pi_t^*$. Hence, the time varying intercept encompass both, changes in equilibrium real interest rate caused by economic fundamentals and inflation target.[11] Horizon i was

9 The policy neutral rate is typically defined as the sum of the real equilibrium rate and inflation target ($r = rr + \pi^*$).

10 The definition of the inflation target varies slightly across IT countries. However, the target is usually mid-term rather than short-term. The target value can also vary over time. This variation has been especially pronounced in emerging countries that implemented IT as a gradual disinflation strategy. By contrast, for the countries studied here, the target value has not changed significantly over time.

11 Clarida et al. (2000, p. 154) note that in absence of further assumptions this approach does not allow identification of economy's equilibrium real rate rr and inflation target π^*. There are various ways how to overcome this identification problem in the literature. Clarida et al. (1998, 2000) assume that rr can be approximate by sample average of real interest rate, Laubach and Williams (2003) use a set of state space models to recover time variation in equilibrium interest rate. Their approach is used by Leight (2008) as well. In this paper, we focus mainly on the evolution of policy parameters β, γ and δ and we leave a detailed decomposition of time-varying intercept for future research.

set to 2.[12] Consequently, the disturbance term ε_t is a combination of forecast errors and is thus orthogonal to all information available at time t (Ω_t).

In line with our previous discussion, the interest rate rule described above will be estimated within a framework that allows time variance of the coefficients. Kim (2006) shows that the conventional time-varying parameter model (the Kalman filter applied to a state-space representation) delivers inconsistent estimates when the explanatory variables are correlated with the disturbance term. Endogeneity arises in forward-looking policy rules based on ex-post data, but it can appear even with real-time data, as discussed before. Kim (2006) proposes an estimator of the time-varying coefficient model with endogenous regressors. A few recent contributions use this framework for estimation of monetary policy rules (Kim and Nelson, 2006, Kim et al. 2006, Kishor, 2008). [13] Following Kim (2006) we can rewrite Eq. (3) as follows:

$$r_t = (1-\rho_t)[\alpha_t+\beta_t(\pi_{t+i})+\gamma_t y_t+\delta_t x_t]+\rho_t r_{t-1}+\varepsilon_t \quad , \tag{4}$$

$$\alpha_t = \alpha_{t-1}+v_{1,t} \ , \ v_{1,t} \sim i.i.d.\, N(0,\sigma^2_{v_1}) \quad , \tag{5}$$

$$\beta_t = \beta_{t-1}+v_{2,t} \ , \ v_{2,t} \sim i.i.d.\, N(0,\sigma^2_{v_2}) \quad , \tag{6}$$

$$\gamma_t = \gamma_{t-1}+v_{3,t} \ , \ v_{3,t} \sim i.i.d.\, N(0,\sigma^2_{v_3}) \quad , \tag{7}$$

$$\delta_t = \delta_{t-1}+v_{4,t} \ , \ v_{4,t} \sim i.i.d.\, N(0,\sigma^2_{v_4}) \quad , \tag{8}$$

$$\rho_t = \rho_{t-1}+v_{5,t} \ , \ v_{5,t} \sim i.i.d.\, N(0,\sigma^2_{v_5}) \quad , \tag{9}$$

$$\pi_{t+i} = Z'_{t-1}\xi+\sigma_\phi \phi_t \ , \ \phi_t \sim i.i.d.\, N(0,1) \quad , \tag{10}$$

$$y_t = Z'_{t-1}\zeta+\sigma_\psi \psi_t \ , \ \psi_t \sim i.i.d.\, N(0,1) \quad , \tag{11}$$

The measurement equation (4) of the state-space representation is the monetary policy rule. The transition equations (5)–(9) describe the time-varying coefficients as a random walk process without drift. Eqs. (10) and (11) track the relationship between the endogenous regressors π_{t+i} (an y_{t+j}) and their instruments, Z_t. The list of instruments, Z_{t-1}, is as follows: π_{t-1}, π_{t-4}, y_{t-1}, y_{t-2}, r_{t-1} and r_t^f (foreign interest rate). Following Kim (2006), we assume that the parameters in Eqs. (10) and (11) are time-varying, too. Contrary to the equation representing the monetary policy rule (4), variances of time-varying coefficients in equations (10) and (11) were not estimated, but calibrated to 0.01. This value assures that coefficients are allowed to vary over time to capture changes in forecasting relationships of expected inflation and output gap, but the estimated paths of coefficients are without large jumps and quick reversals.[14] Next, the correlation between the standardized residuals φ_t, ψ_t and ε_t is $\kappa_{\varphi,\varepsilon}$ and $\kappa_{\psi,\varepsilon}$, respectively (note that σ_φ and σ_ψ are standard errors of φ_t and ψ_t, respectively). The consistent estimates of the coefficients in Eq. (4) are obtained

12 Although the targeting horizon of central banks is usually longer (4–8 quarters), we prefer to proxy inflation expectations by inflation in t+2 for the following reasons. First, the endogeneity correction requires a strong correlation between the endogenous regressor and its instruments. Second, the prediction error logically increases at longer horizons. Third, the countries we analyze did not apply inflation targeting during the whole estimation period. Consequently, it is preferable owing to data limitations to keep only two inflation leads rather than four or six.

13 Note, however, that two of these contributions are, to our knowledge, unpublished as yet.

14 Interestingly, predicted inflation seems to be very similar to long-term inflation expectations of central banks of the United Kingdom, Sweden and New Zealand shown in Kuttner (2004, p. 97, Figure 1).

in two steps. In the first step, we estimate equations (10) and (11) and save the standardized residuals φ_t and ψ_t. In the second step, we estimate Eq. (12) below along with Eqs. (5)–(9). Note that (12) now includes bias correction terms[15], i.e. the (standardized) residuals from Eqs. (10) and (11), to address the aforementioned endogeneity of the regressors. Consequently, the estimated parameters in Eq. (12) are consistent, as instruments u_t is uncorrelated with the regressors.

$$\begin{aligned} r_t &= (1-\rho_t)[\alpha_t+\beta_t(\pi_{t+i})+\gamma_t y_t+\delta_t x_t]+\rho_t r_{t-1}+\kappa_{\phi,\varepsilon}\sigma_{\phi,\varepsilon}\phi_t+\kappa_{\psi,\varepsilon}\sigma_{\psi,\varepsilon}\psi_t+u_t\ ,\\ u_t &\sim N(0,(1-\kappa^2_{\phi,\varepsilon}+\kappa^2_{\xi,\varepsilon})\sigma^2_{\varepsilon,t}) \end{aligned} \tag{12}$$

The standard framework for estimation of Eqs. (10), (11) and (12) is the maximum likelihood estimator via the Kalman filter (Kim, 2006). However, there are several difficulties with the estimation of the Kalman filter (and Kalman smoother) in applied work. First, if the variables are nonstationary, the results often depend on the proper choice of the initial values, but those values are not known in advance.[16] The problem with the initial conditions is larger if one-sided estimates are used, as illustrated in Leigh (2008) on estimates of the time-varying natural rate of interest in the U.S. Applying the Kalman smoother alleviates the issue and, for different initial values, the differences in the estimates at the beginning of the sample decrease sharply. Second, the log likelihood function is highly nonlinear and in some cases the optimization algorithm fails to minimize the negative of the log likelihood for several reasons (either it can fail to calculate the Hessian matrix throughout the iteration process or, when the likelihood function is approximated to facilitate the computations, the covariance matrix of the observation vector can become singular for the provided starting values).

In this paper, we adopt the "varying coefficients" (VC) method (Schlicht and Ludsteck, 2006). The VC method generalizes the standard ordinary least squares approach. In fact, instead of minimizing the sum of the squares of the residuals , $\sum_{t=1}^{T} u_t^2$ it uses minimization of the weighted sum of the squares:

$$\sum_{t=1}^{T} u_t^2+\theta_1\sum_{t=1}^{T} v_{1,t}^2+\theta_2\sum_{t=1}^{T} v_{2,t}^2+\ldots+\theta_n\sum_{t=1}^{T} v_{n,t}^2 \tag{13}$$

where the weights θ_i are the inverse variance ratios of the regression residuals u_t and the shocks in time-varying coefficients v_t , that is $\theta_i=\sigma^2/\sigma_i^2$. Hence it balances the fit of the model and the parameter stability.[17.] Additionally, the time averages of the regression coefficients estimated by such weighted least squares estimator are identical to the GLS estimates of the corresponding

15 Obviously, if the correction terms are statistically significant, it shows that endogeneity matters. Similarly to Kim and Nelson (2006) and Horváth (2009), we find that these terms are significant: in our sample the endogeneity correction for inflation is significant for the United Kingdom, Canada and Sweden at 5% level, and for the GDP gap it is significant for Canada (see table A.1 in the Appendix).

16 Although there are a number of formal procedures for initialization of the Kalman filter in such cases (for example see Koopman et al., 1999), fundamental uncertainty about their values remains.

17 It should be noted that throughout our computations we did not have to solve problems with convergence of the moment estimator, as it was almost always able to find equilibrium. Computational details of the VC method are described in the Appendix. Originally, Schlicht and Ludsteck (2006) start with a derivation of the maximum likelihood estimator of parameters a based on the idea of orthogonal parameterization, which is described in the Appendix. Then they prove that the weighted least squares estimator is identical to the maximum likelihood estimator and also that the likelihood estimator is identical to the moment estimator for very large samples.

regression with fixed coefficients, that is, $\frac{1}{T}\sum_{t=1}^{T}\hat{a}_t = \hat{a}_{GLS}$.

The VC method has a number of advantages. First, it does not require initial conditions even for non-stationary variables prior to the estimation procedure. Instead, both the variance ratios and the coefficients are estimated simultaneously. Second, the property of the estimator that the time averages of the estimated time-varying coefficients are equal to their time-invariant counterparts permits easier interpretation of the results by comparison with time-invariant results. The features of the VC method make it feasible for our analysis: we deal with a time-varying model where the coefficients are assumed to follow a random walk, there is no *a priori* information about the initial values and the time series are rather short.[18]

Furthermore, Schlicht and Ludsteck (2006) compare the results from the VC method and from the Kalman filter, showing that both estimators give very similar results given the assumption that the Kalman filter is initialized with the correct initial conditions. Yet in this case, the VC estimator has a slightly lower mean squared error and this difference is more pronounced for small samples.[19]

We assume that the variance of the disturbance term in Eq. (12) is not time-varying. Nevertheless, there is an ongoing discussion about to what extent changes in the macroeconomic environment are driven by changes in the variance of the disturbance term (i.e. exogenous changes in the economic system) vis-à-vis the variance in the coefficients of the monetary policy rule (see, for example, Benati and Surico, 2008, Canova and Gambetti, 2008, or Sims and Zha, 2006).

One can also think about Eqs. (4), (10) and (11) in terms of the New Keynesian model, with Eqs. (10) and (11) representing the Phillips and IS curves. It should be noted that our framework is in general less restrictive and imposes less structure than the full-blown New Keynesian model.

We expect β_t to be positive, as the central bank is likely to react to an increase in expected inflation by increasing its policy rate. In particular, β_t should be greater than one in the long-run solution of Eq. (4) if monetary policy is stabilizing. The development of β_t over time may be driven by a number of factors, such as changes in monetary policy regime or institutional constraints (Adam et al., 2005). The effect of the adoption of inflation targeting on β_t is ambiguous. As put forward by Kuttner and Posen (1999), β_t can both increase and decrease. They show that under a conservative central bank the response of short-term interest rates is greater than under discretion or

18 The number of observations differs across the countries, ranging from 103 to 144. In the case of Kalman filter/smoother we can utilize the whole sample if we opt for initial conditions equal to the full sample OLS estimated values (recommended for stationary systems). Another approach derives the initial conditions related directly to the beginning of the sample from the first subset of available observations and the Kalman filter is performed on the latter part of the sample. Kim and Nelson (2006) adopted this approach and used the first 40 observations for the initialization. The estimation of the second step is carried out by Schlicht's VC package, which uses the moment estimator.

19 For comparison, we estimated equation (12) using the conventional Kalman filter/smoother in the GROCER software using the function tvp (Dubois-Michaux, 2009). We parameterized the model with initial conditions taken from the OLS estimates of the parameters on the full sample and the initial forecast error covariance matrix set to 0. The matrix of the residuals of the time-varying coefficients is assumed to be diagonal as in the VC method. The results for Kalman smoother were very similar to those obtained from the VC method, with the estimated variances being the same in both methods. The only country where the estimated variance was different, was Sweden, with a lower variance in smoothing parameter ϱ and higher a variance in β. Still, the results were consistent with ours. These results are available upon request.

the optimal state-contingent rule (inflation targeting),[20] while the strength of the response under inflation targeting as compared to discretion depends on the credibility of the regime. Credible monetary policy does not have to react so strongly to inflation surprises, as inflation expectations are likely to remain anchored. Sekine and Teranishi (2008) provide a new Keynesian model that reaches to the same conclusions. Siklos and Weymark (2009) estimate that inflation targeting in Australia, Canada and New Zealand reduced the magnitude of the interest rate changes to needed to maintain a low inflation environment.

Similarly, ϱ_t, a measure of interest rate smoothing, is expected to be positive with values between zero and one. Many time-invariant estimates of monetary policy rules find the value of this parameter to be about 0.7–0.9, implying a substantial degree of interest rate smoothing. Rudebusch (2006) claims that such figures are clearly overestimated in the face of very low interest rate forecastability in the term structure of interest rates. On the contrary, the time-varying model in principle enables some variables to affect interest rate setting in one period but not in another, and is less prone to autocorrelated shocks.

Next, the effect of the output gap, γ_t, on interest rates is expected to be positive or insignificant. In the first case, the central bank may have an explicit concern for real activity or understand the output gap as a useful predictor of future inflation. In the latter case, the insignificant coefficient may suggest that the central bank is primarily focused on inflation and does not consider the output gap to be important in delivering low inflation.

There is a debate in literature about whether other variables should be included in the monetary policy rule. This is especially appealing for small open economies, which may be concerned with exchange rate fluctuations as well as the evolution of foreign interest rates. Taylor (2001) puts forward that even if the exchange rate or foreign interest rates are not explicitly included in the policy rule, they still remain present implicitly, as the exchange rate influences the inflation forecast, to which the inflation-targeting central bank is likely to react. It is also worth emphasizing that significance of the exchange rate or foreign interest rates does not necessarily mean that the central bank targets some particular values of these variables, but rather that the bank considers foreign developments to be important for its inflation forecast. On the other hand, empirical studies often favor the inclusion of these variables in the estimated policy rule. Having these considerations in mind, we decided to include the exchange rate and foreign interest rates (of dominant economies with respect to each analysed country), too, in order to assess whether these two variables carry any additional information for understanding the interest rate setting process in our sample countries.

2.3.2 The dataset

We use quarterly data. The sample period varies from country to country owing to data availability (the UK 1975:1Q–2007:4Q, Australia 1975:1Q–2007:4Q, Canada 1975:1Q–2007:4Q, New Zealand 1981:1Q–2007:4Q, Sweden 1982:2Q–2007:3Q), but on average the time coverage is about three decades.

Following Clarida, Galí and Gertler (1998), the dependent variable is the short-term interest rate, which is typically closely linked to the monetary policy rate. The reason for choosing the short-term interest rate rather than the monetary policy rate is the fact that the monetary policy rate and the change therein are censored (Podpiera, 2008). Therefore, the dependent variables capturing the

20 See King (1997) on how inflation targeting allows one to come close to the optimal state-contingent rule.

policy rate are the discount rate (3-month Treasury bills) for the UK, the interbank 3-month interest rate for Australia, the 3-month Treasury bills rate for Canada and the interbank 3-month interest rate for New Zealand and Sweden. We choose the interest rate so as to be closely linked to monetary policy, but also to be available for a sufficiently long period. The foreign interest rate is the German 3-month Euribor for the UK and Sweden and the U.S. 3-month interbank interest rate for Australia, NZ, and Canada.

The inflation is measured as the year-on-year change in the CPI, except for the UK, where we use the RPIX (the retail price index excluding mortgage interest payments), and the NZ, where we use the CPIX (the CPI without interest payments).[21] The output gap is taken as reported in the OECD Economic Outlook (the production function method based on the NAWRU – the non-accelerating wages rate of unemployment).[22] The exchange rate is measured by the chain-linked nominal effective exchange rate (NEER). For the regressions we use the deviation of the index from the HP trend (first differences and yearly changes were used for a robustness check).[23]

2.4 Results

First, this section presents the country-specific estimates of the time-varying monetary policy rules in sub-sections 2.4.1 – 2.4.5. The sub-section 2.4.6 contains the estimates of time-varying inflation persistence and the sub-section 2.4.7 provides the summary of the main policy-relevant findings.

2.4.1 United Kingdom

Our results show that the BoE significantly increased its response to inflation since late 1970's till mid-eighties. This overlaps with the Thatcher government and its major priority being the inflation control. The overall decline of the response since 1985 can be related to the dismissal of medium-term financial strategy (adopted in 1979). We find that the response of interest rates on inflation was gradually decreasing during the 1990s in spite of the introduction of the IT. Although this finding can seem at the first sight contra-intuitive, it is important to keep in mind that, unlike in some emerging countries, the IT was not implemented in the UK as a strong anti-inflationist strategy. The inflation was already contained in the 1980's and very benign inflation enviroment was also supported by declining prices of raw materials on the world markets. This corroborates with the findings of Kuttner and Posen (1999) and Sekine and Teranishi (2008) who show that inflation targeting can be associated with less aggressive monetary policy.

The effect of the output gap is estimated as positive (albeit the confidence intervals are rather large, probably reflecting the fact that the gap is an unobserved variable and calculated ex post) and does not vary substantially over time. The interest rate smoothing parameter is found to have values between 0 and 0.2, which is much lower than typically reported by time-invariant estimates of monetary policy rules (Clarida et al., 1998, 2000). Our estimates seem to be reasonable in the face

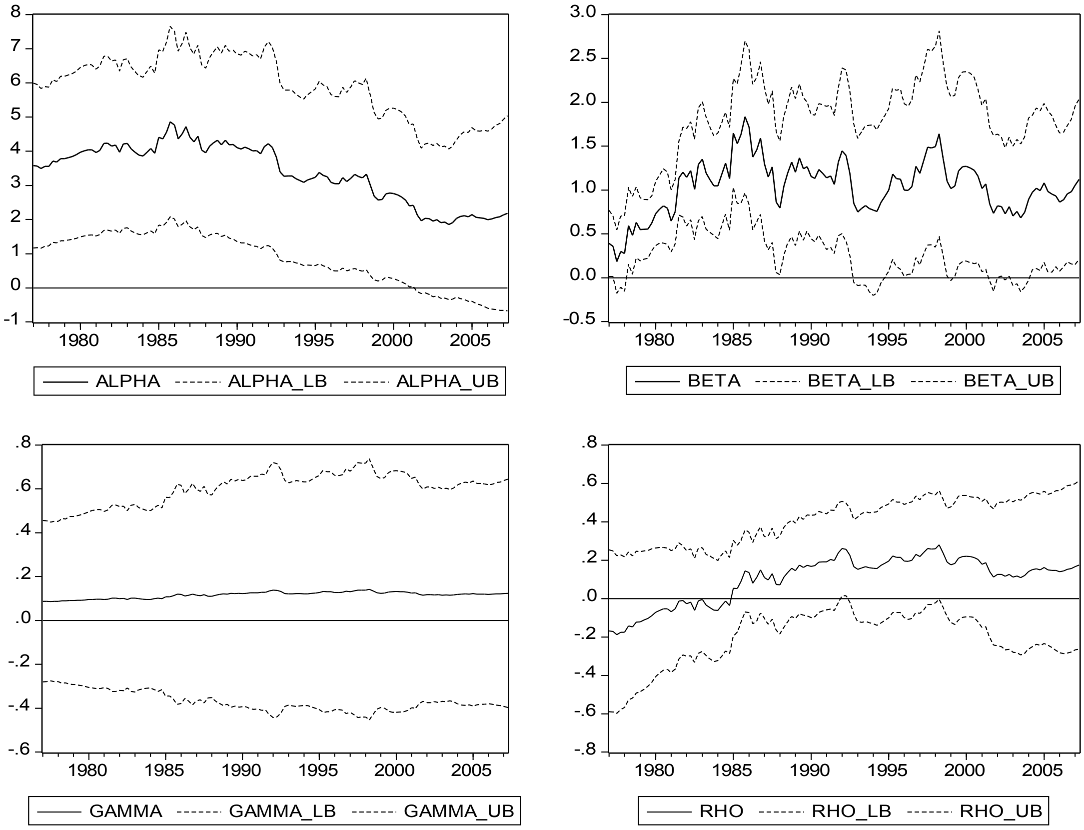
21 We use year-on-year data, as the inflation target is also defined on a year-on-year basis.

22 There is no agreement on what is the best method for extraction of the unobserved output gap (Billmeier, 2009). We prefer to use the output gaps obtained by the OECD by means of the production function approach because they are based on a substantially richer information set than simple statistical detrending. Somewhat surprisingly, the OECD output gap and the simple HP gap evolve very closely.

23 It is rather puzzling what measure of exchange rate movements should be used to test potential central banks response. Given that the IT central banks do not declare exchange rate targets, we must rely on some measure of exchange rate missalightment.. We use three different alternatives present in literature (see e.g. Clarida et al., 1998 or Lubik and Schorfheide, 2007). All time series used are shown in Appendix 2.

of the recent critique by Rudebusch (2006). Finally, the intercept can be in this basic model interpreted as policy neutral (nominal) interest rate. We can see that it steadily declined over time, which complies with the low inflation environment in the 1990s that prevailed in the U.K.

Figure 1 – Time-varying response coefficients in baseline (closed economy) policy rule, UK



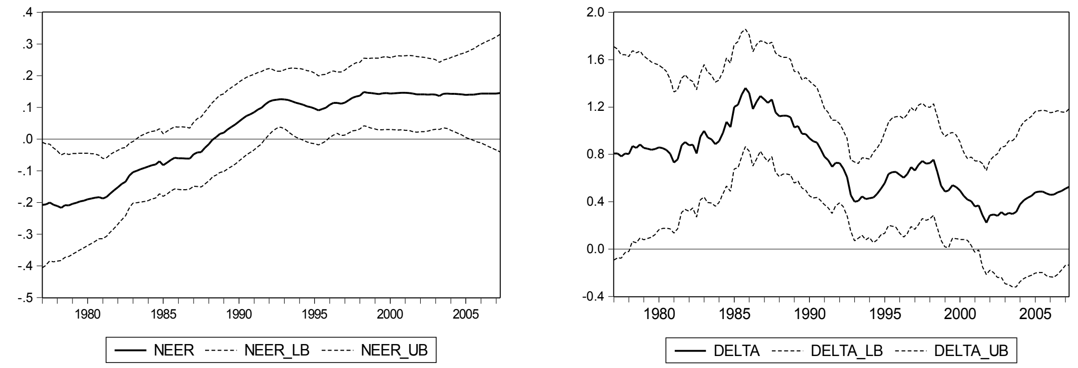
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the time-varying intercept. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

The results of our augmented model show that the monetary policy of the BoE was influenced by external factors, although their importance was greater in the 1980s than recently.[24] In particular, we find evidence that the BoE decreased its policy rate as the nominal effective exchange rate (NEER) strengthened during the 1980s even before the pound officially joined the ERM (1990). Once the UK abandoned the ERM and introduced IT, the response to the exchange rate turned slightly positive and practically invariant. These results are consistent across different transformation of the exchange rate such as deviation from HP trend, first differences and year-on-year change. Obviously, it considered the exchange rate indirectly, as exchange rate fluctuations influence the inflation forecast (for more on this see Taylor, 2001). The same reasoning applies to the response to the foreign interest rate (Euribor). It was particularly strong during the 1980s and subsequently its importance somewhat declined. Our results show little support for the hypothesis that the monetary policy of the BoE follows that of the ECB, as the estimated response of the coefficient declined and the confidence intervals widened after the launch of the euro.

24 In what follows, we present the evolution of the coefficients for the response to the exchange rate and foreign interest rates; the other coefficients remain largely unchanged and are not reported for the sake of brevity.

There are two directly comparable studies to this paper. Kishor (2008) obtains for the UK results similar to ours in spite of using monthly data known to have slightly different dynamics. He finds that the anti-inflation stance peaked in the mid-1980s and tended to decline from then onwards in spite of the adoption of IT. Similarly, his finding that the response to foreign interest rate significantly declined since the ERM crises is complementary to our result that the BoE was giving much less consideration to exchange rate evolution (NEER gap).

Figure 2 – Time-varying response coefficients in augmented (open economy) policy rule, UK



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

Trecroci and Vassalli (2010), who, unlike Kishor (2008) and this paper, do not correct for endogeneity in the time-varying model, come to the opposite conclusion that the BoE's response to inflation increased over time. Yet, some counterintuitive results of their study point to the possibility of endogeneity bias. First, the interest rate smoothing parameter takes on significantly negative values from 1980 till 1995. This would imply not only that policy was not inertial, but also that there was actually a negative correlation between the present and past interest rate, which is inconsistent even in the face of a simple visual inspection of the interest rate series. Second, their coefficient for the foreign (German) interest rate peaks in 1990 and is *de facto* invariant since then, which the authors interpret as implicit exchange-rate targeting. This finding is doubtful given the pound's exit from the ERM and the implementation of IT from 1992 onwards. In fact, British and German short-term rates, which were almost at par in 1992, diverged and the interbank interest rate in the UK exceeded the German one by almost 4% on the eve of euro adoption.

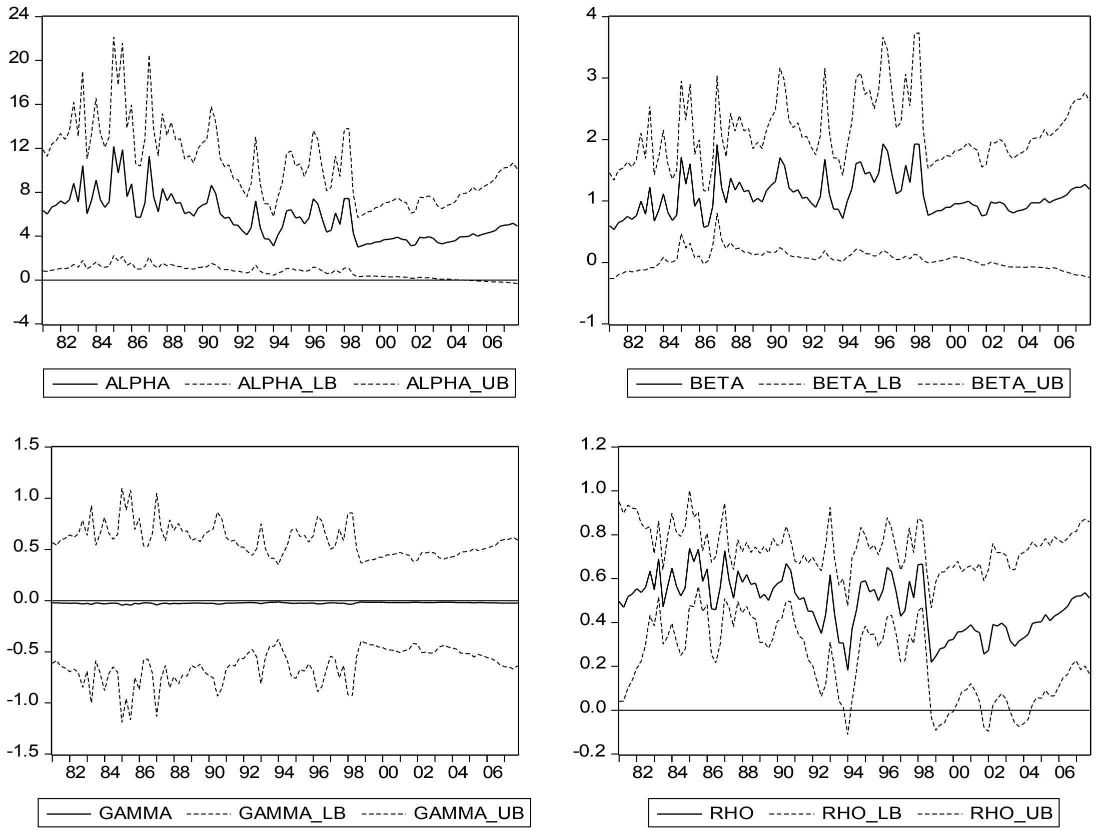
2.4.2 New Zealand

The inflation targeting was introduced in New Zealand as the first country in the world by the Federal Bank Act signed in March 1990.[25] Our results indicate that the response of the RBNZ to expected inflation was very close to unity throughout the sample period (1981–2007). In fact, it is clearly visible that the interest rate and inflation series move together very closely. Therefore, in Figure 3 we can see that the official introduction of IT does not seem to have engendered a

25 Huang et al. (2001) argue that this policy was in effect since the end of 1988, when the RBNZ abandoned both monetary and exchange rate targeting. They also point to a specific feature of RBNZ monetary policy that could be referred to as "Open Mouth Operations". Between 1989 and 1999 the RBNZ specified a 90-day bank bill rate consistent with price stability and threatened to use quantitative controls to achieve the desired market rate if it deviated from the target. Therefore, the RBNZ did not control this interest rate permanently and directly.

significant change in interest rate setting (if anything there is very slight decrease of the response coefficient on inflation after 1998). Unlike in the UK, the response coefficient does not decrease substantially right after the IT adoption. This may be related to the fact that at the time IT was introduced in New Zealand the inflation rate was not far from double-digit values. Therefore, this policy was implemented in a different context than, say, in the UK, where single-digit inflation had already been achieved during the 1980s. This result, together with the estimated insignificant response to the output gap, is consistent with the findings of time-invariant studies (Huang et al., 2001; Plantier and Scrimgeour, 2002) that the RBNZ applied a rather strict version of inflation targeting. On the other hand, in 1999 the objectives of the RBNZ changed. Since then, the RBNZ should have avoided unnecessary instability in output, interest rates and exchange rate and in consequence, changes in interest rate are actually less frequent after. This change in monetary policy is clearly reflected by smoother coefficients in comparison to the previous periods.

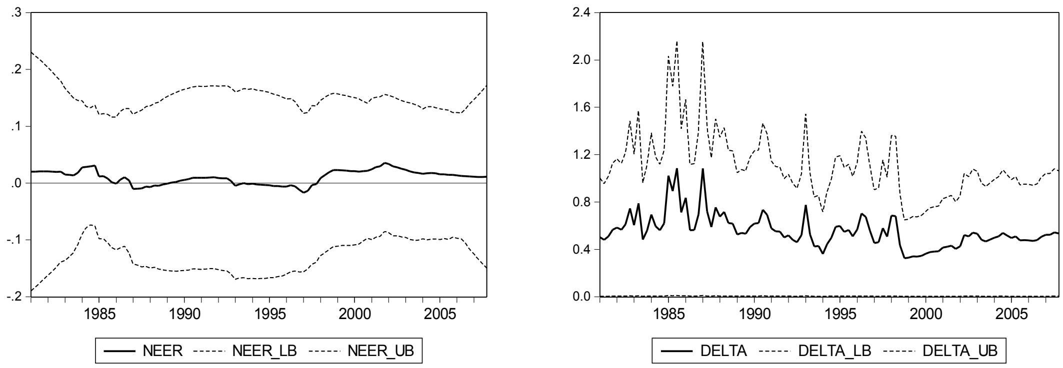
Figure 3 – Time-varying response coefficients in closed economy policy rule, New Zealand



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the time-varying intercept. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

When we estimate the augmented model for New Zealand, we do not find any indication that the exchange rate was ever considered by the RBNZ for interest rate setting. This finding is again confirmed when we use exchange rate in differences and yearly changes. Our evidence is consistent with Ftiti (2008) who in a time-invariant model rejects the hypothesis that the RBNZ responded to the exchange rate. On the contrary, we find some evidence in favor of consideration of the foreign interest rate, although its response coefficient generally decreased after the launch of IT.

Figure 4 – Time-varying response coefficients in open economy policy rule, New Zealand



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

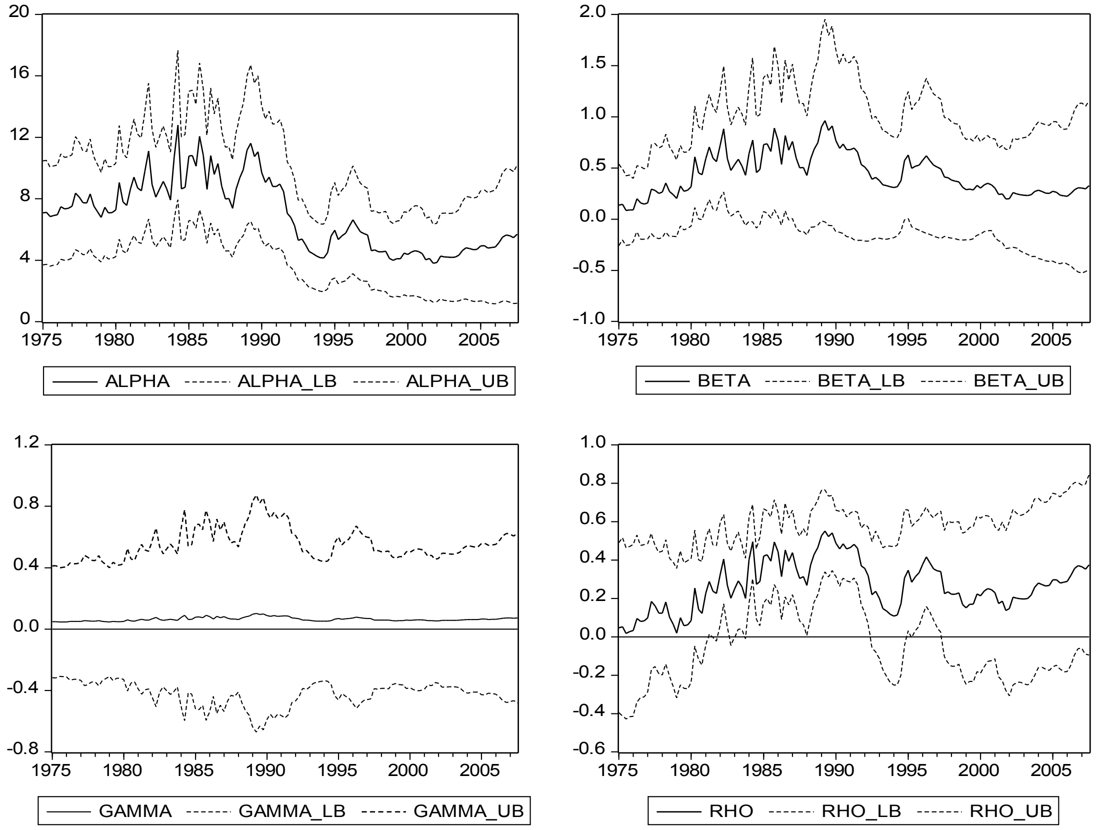
2.4.3 Australia

Our results for Australia are available in Figures 5 and 6. The response of the interest rate to inflation is strongest in the 1980s, which is very similar to the UK experience. This period was characterized by inflation rates of around 10% and central bankers had to be quite aggressive in interest rate setting in order to break the record of high inflation deeply ingrained in public expectations. Neither monetary targeting (employed until 1984) nor the checklist approach (1985–1990) seemed to be successful in this regard. The fluctuation of the inflation response coefficient points to the discretionary nature of policy decisions (making this finding consistent with Leu and Sheen, 2006). The response coefficient peaks in 1990 on the eve of IT but declines after the adoption of this regime. It is again arguable whether it was the credibility of this regime that anchored inflation expectations and allowed the RBA to behave less aggressively. The original inflation decline may also have been related to the world recession in the early 1990s. Our results dispute the finding of De Brouwer and Gordon (2005) that the inflation response of the RBA increased as a result of the launch of inflation targeting.

As for other countries, the time-varying intercept and arguably also the policy neutral rate decline in the 1990s, reflecting the global low inflation environment. The output gap is not found to be significant and the estimated interest rate smoothing is again rather low.

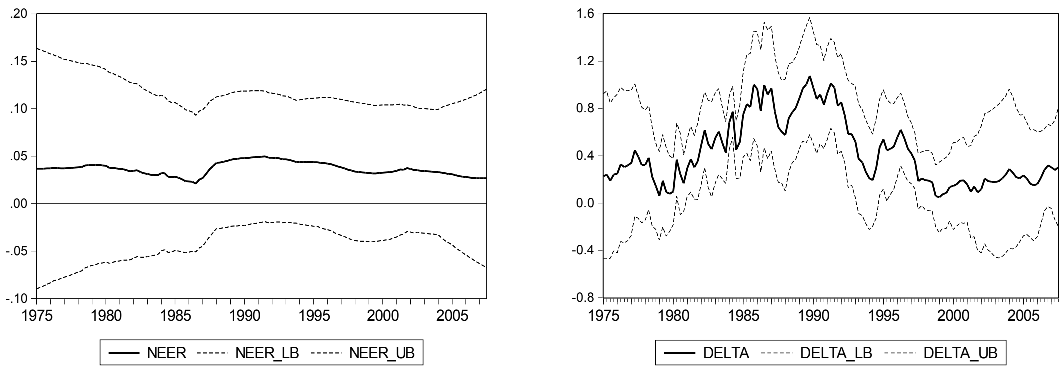
We find that the exchange rate does not have a significant effect on the short-term interest rate (besides the NEER we use also the trade-weighted index – TWI, which is an exchange rate measure reported and often referred to by the RBA). This finding is again consistent across different exchange rate transformations. The foreign interest rate parameter is estimated as being always positive, although it is significant only in the 1990s and its importance fades after IT was introduced. After 2001, Australian and U.S. interest rates diverge and the response coefficient approaches zero. This may be related to idiosyncratic developments in the U.S. when the Fed lowered the interest rate so as to face the fear of recession following the September 2001 terrorist attacks.

Figure 5 – Time-varying response coefficients in baseline (closed economy) policy rule, Australia



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the time-varying intercept. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

Figure 6 – Time-varying response coefficients in augmented (open economy) policy rule, Australia



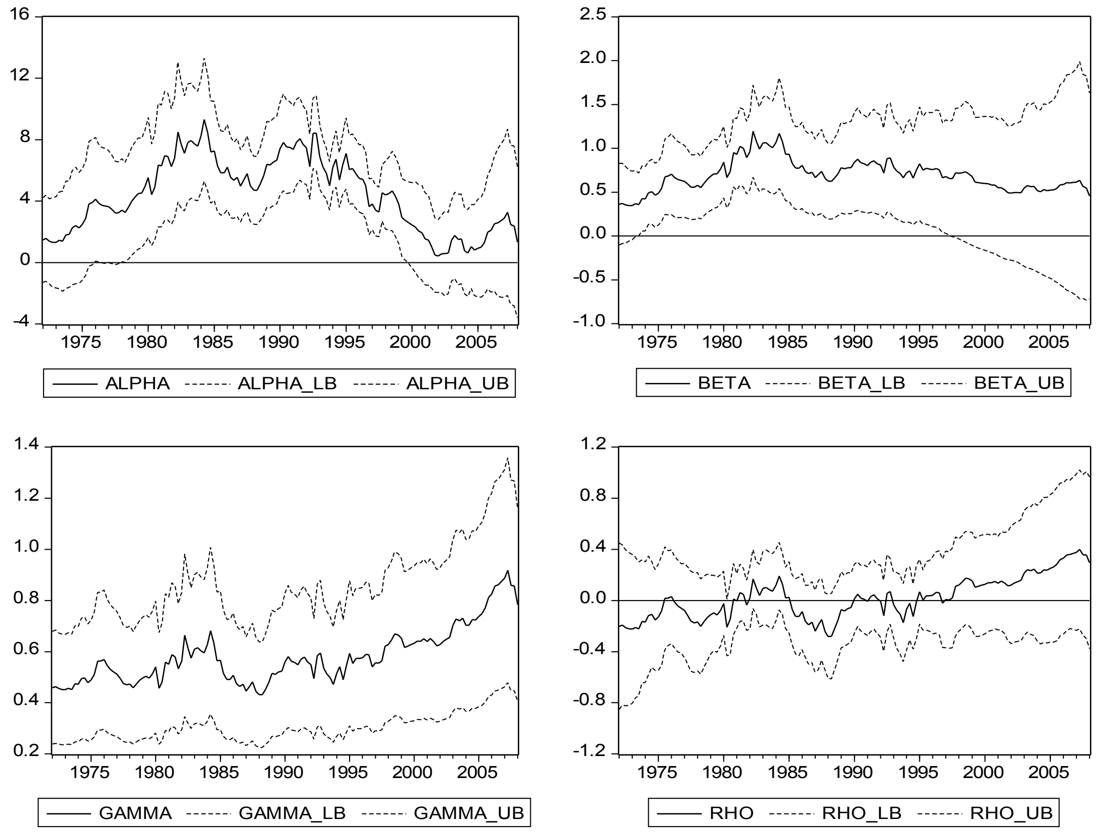
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

2.4.4 Canada

The monetary policy rule estimates for Canada are presented in Figures 7 and 8. The response of the interest rate to inflation peaks in the mid-1990s, which was a period characterized by relatively high inflation rates, which unquestionably drove the rather aggressive policy of the BoC similarly as in the UK and Australia. It is arguable whether the original inflation rate was a consequence of the accommodative policy of monetary targeting applied between 1978 and 1982 (see Figure 7). Since mid-1980s the inflation response coefficient has been steadily declining and the IT adoption in 1993 did not change its course. Almost neglibigle inflation rates in the last decade drove very significant widening of the confidence intervals.[26]

The response to the output gap is possitive and often statistically significant, confirming the long-term preference of the BoC for smoothing economic fluctuations. The intensity of the response is unique among the IT countries in our sample. The interest rate smoothing is almost negligible and the time-varying constant terms shows, as in other countries, decreasing pattern since the early 1990s.

Figure 7 – Time-varying response coefficients in baseline (closed economy) policy rule, Canada

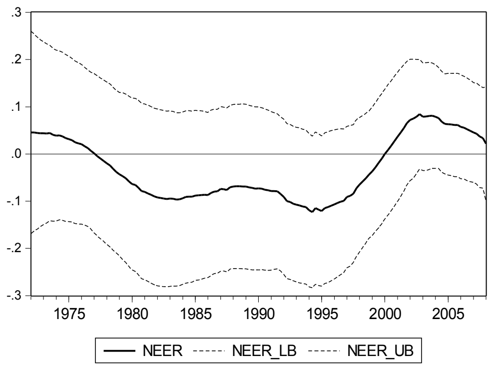


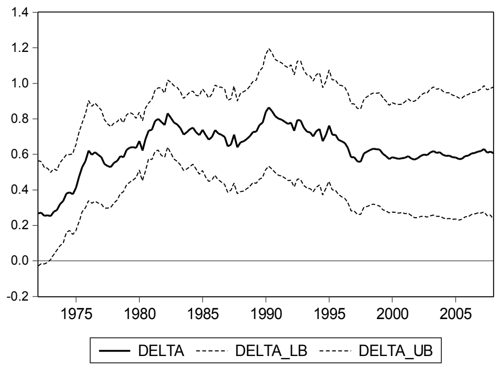
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the time-varying intercept. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

26 The BoC also reported the monetary condition index (MCI), a compound of the policy instrument (the interest rate) and the exchange rate. The MCI accompanies the proposal of Ball (1999) to target long-term inflation, i.e. the inflation rate adjusted for the transitory effect of the exchange rate on import prices. However, there is no indication that the BoC actually ever used the MCI for practical policy making, and it ceased to publish it in 2006.

The dependence of Canadian monetary policy on external factors, in particular developments in the U.S., is confirmed in the model augmented by the foreign interest rate. The response to the U.S. interest rate dynamics is substantial for the whole period of analysis until the end of the sample. The response to the exchange rate is mostly negative (decrease of policy rate when the exchange rate is appreciating), though mostly insignificant. This patterns is not altered when other exchange rate transformations are used.

Figure 8 – Time-varying response coefficients in augmented (open economy) policy rule, Canada





Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

2.4.5 Sweden

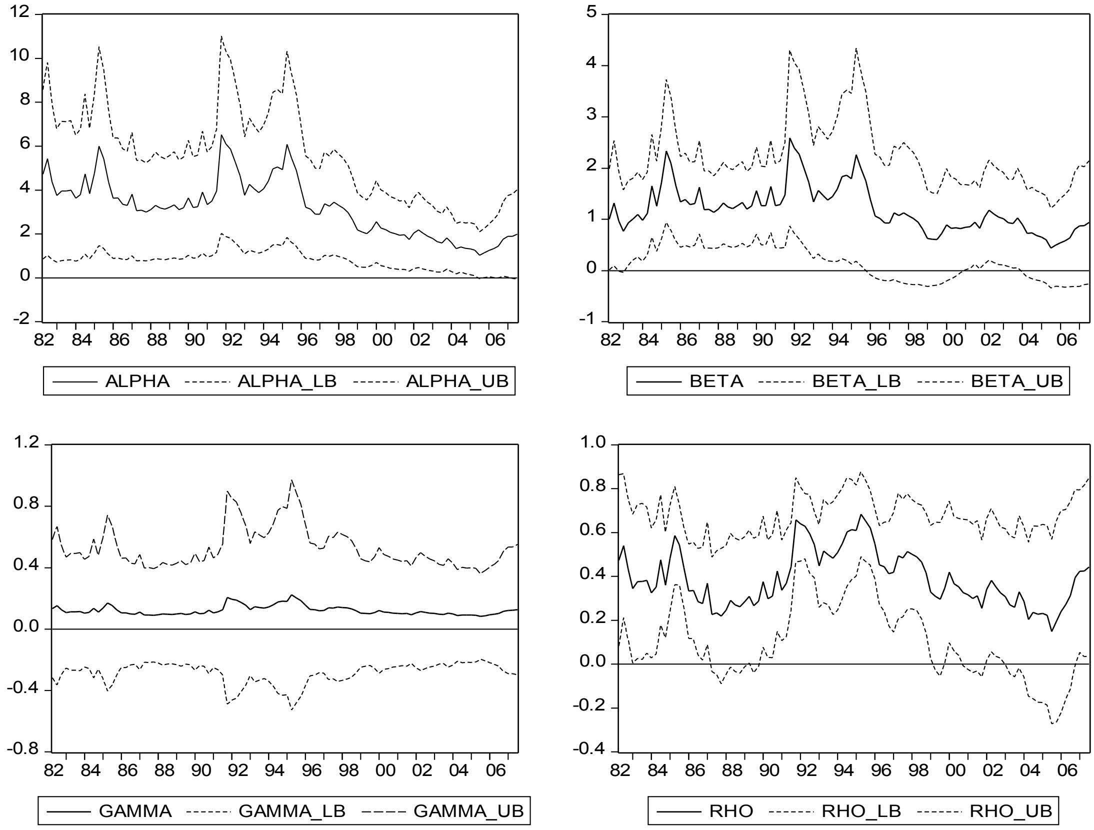
Our results suggest that the response of interest rates to inflation was stronger before and at the beginning of IT. This is in line with Berg et al. (2004), who argue that the introductory phase of IT in Sweden was characterized by building the credibility of the new regime. Sveriges Riksbank seems to have disregarded, as most other central banks considered here, the output gap. The decline of the time-varying intercept reflects the low inflation environment prevailing in Sweden from the mid-1990s onwards. The time-varying coefficient on interest rate smoothing is estimated to be somewhat larger in Sweden, alike in Australia, than in the other three countries. This suggests that Sveriges Riksbank is likely to smooth its interest rates to a greater degree.[27]

The external factors have a prominent role for the determination of Swedish monetary policy. In particular, the coefficient on the foreign interest rate (Eurolibor) is sizeable throughout the whole sample period, which is rather interesting given that Swedish monetary policy has not officially been subject to any external constraint (at least since the krona's exit from the ERM in 1992). The NEER response coefficient is mostly positive, although with wide confidence intervals. When we use first differences and year-on-year changes of the exchange rate we obtain very similar results.

Overall, our results are consistent with the surveyed time-invariant studies emphasizing the predominant role of the inflation forecast (Jansson and Vredin, 2003) as well as more cautious policy decisions leading to more policy inertia during periods of macroeconomic instability such as the ERM crisis (Berg at al., 2004).

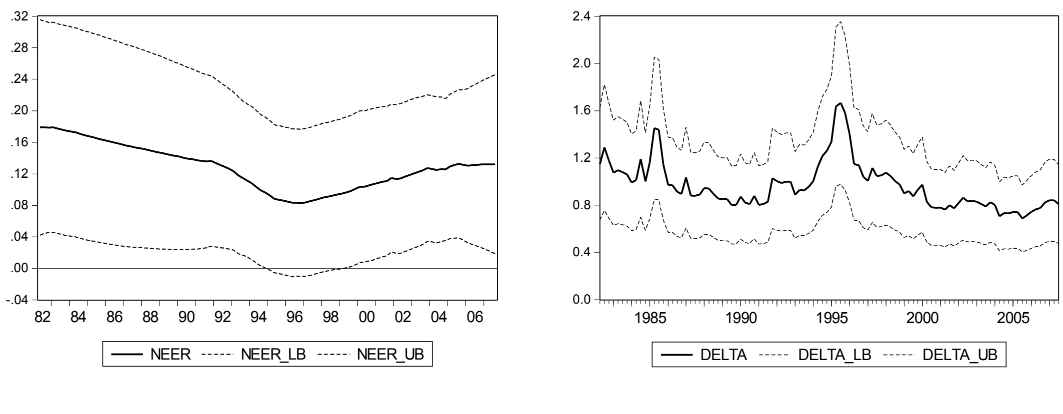
27 At the time of the ERM crisis (September 1992), the Swedish krona started to depreciate. The Sveriges Riksbank tried (unsuccessfully) to maintain the previous exchange rate and massively increased the short-term interest rate. Consequently, we have included a time dummy in Q3 1992.

Figure 9 – Time-varying response coefficients in baseline (closed economy) policy rule, Sweden



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the time-varying intercept. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

Figure 10 – Time-varying response coefficients in augmented (open economy) policy rule, Sweden



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

2.4.6 Inflation Targeting and Inflation Persistence

We have related our finding that the inflation response coefficient often falls after the adoption of IT to the hypothesis that this monetary framework has a positive effect on the inflation expectations of economic agents. If expected inflation is low, monetary policy need not be as aggressive as under a discretionary regime in order to achieve price stability. This argument is in line with recent studies on inflation dynamics (Benati, 2008; Zhang et al., 2008) claiming that under a credible policy regime (such as IT), inflation persistence (the dependence of current inflation on past values) fades away.

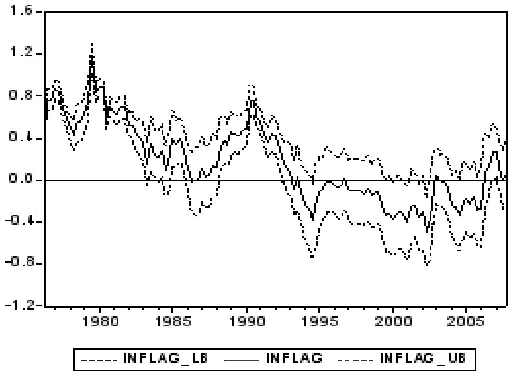
As to shed some light on this issue, we used our estimation framework and fitted the AR(1) model with drift to inflation series, allowing the coefficient on lagged inflation as well as the constant to be time-varying. Our results (as reported in Figure 11) indicate that inflation persistence decreased over time for all the countries. Moreover, it is notable that the persistence fell especially during the 1990s as IT was introduced. This finding is confirmed when, in the spirit of the backward-looking Phillips curve, we include the lagged output gap as a forcing variable.[28] In addition, the results reported in Figure 11 clearly indicate that the moment estimator applied to the time-varying coefficient approach is able to trace also periods when the estimated coefficient is subject to sudden switches rather than smooth transition (see the UK after the adoption of inflation targeting).

In general, our results are broadly consistent with those of Benati (2008), who performs sub-sample analysis under different policy regimes. Unlike Benati (2008), our approach does not need to impose breaks in the inflation process at any particular date, but simply observes whether and when such breaks occur. Our findings do not exclude the possibility that inflation persistence decreased because of other factors (the "good luck" hypothesis), but the temporal coincidence between the introduction of IT and the significant decrease of inflation persistence in several countries make a case for the "good policy" hypothesis. Taking the example of the UK, we can see that the inflation (rate) moderation goes back to the 1980s, when we still observe rather high inflation persistence in spite of very aggressive anti-inflationary (yet discretionary) policy. Unfortunately, as we do not have a full structural model we cannot tell much about the nature of shocks in the pre- and post-IT period as done in VAR studies on the Great Moderation. Using standard tests of structural stability, we can clearly reject structural stability of the inflation process defined by AR(1) in the pre- and post-IT period.

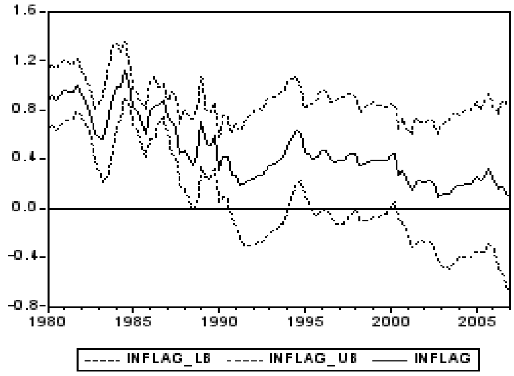
28 See Hondroyiannis et al. (2009) on the estimation of the Phillips curve within a somewhat different time-varying framework.

Figure 11 – Time-varying response coefficients in AR(1) model for inflation

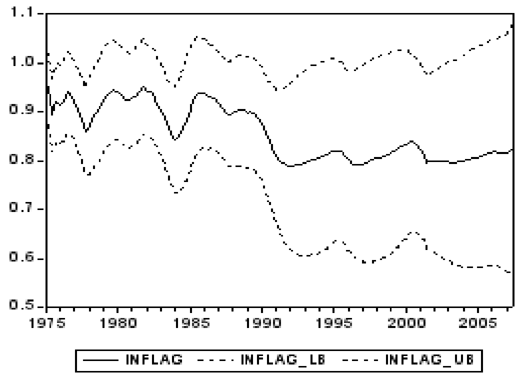
United Kingdom



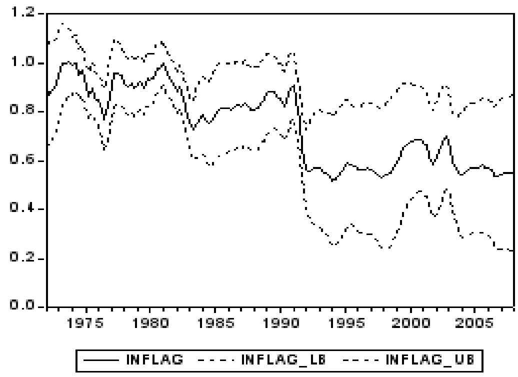
New Zealand



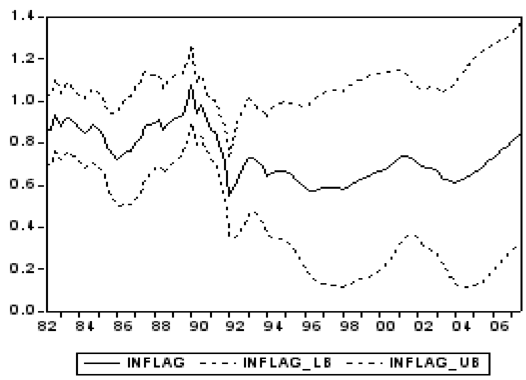
Australia



Canada



Sweden



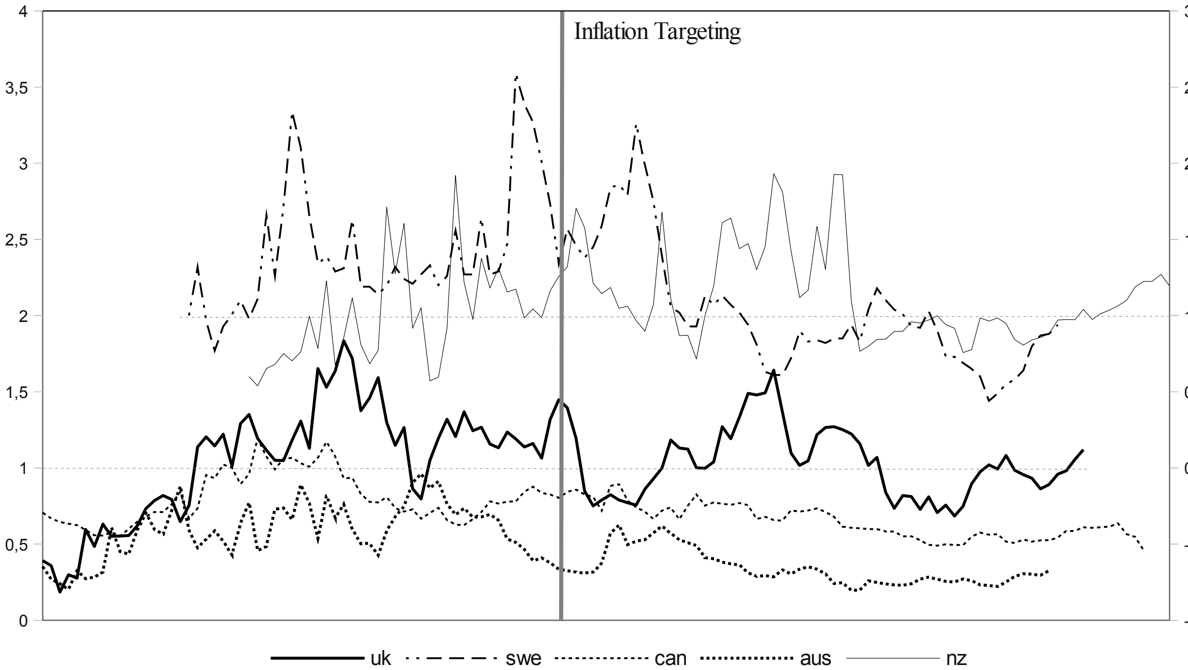
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

2.4.7 Monetary Policy Rules – Wrap-Up of Main Policy Findings

This sub-section summarizes the main policy-relevant findings of this paper. We focus on the following four issues: 1) monetary policy aggressiveness and and inflation targeting, 2) monetary policy aggressiveness and the inflation rate, 3) interest rate smoothing and 4) inflation persistence and inflation targeting.

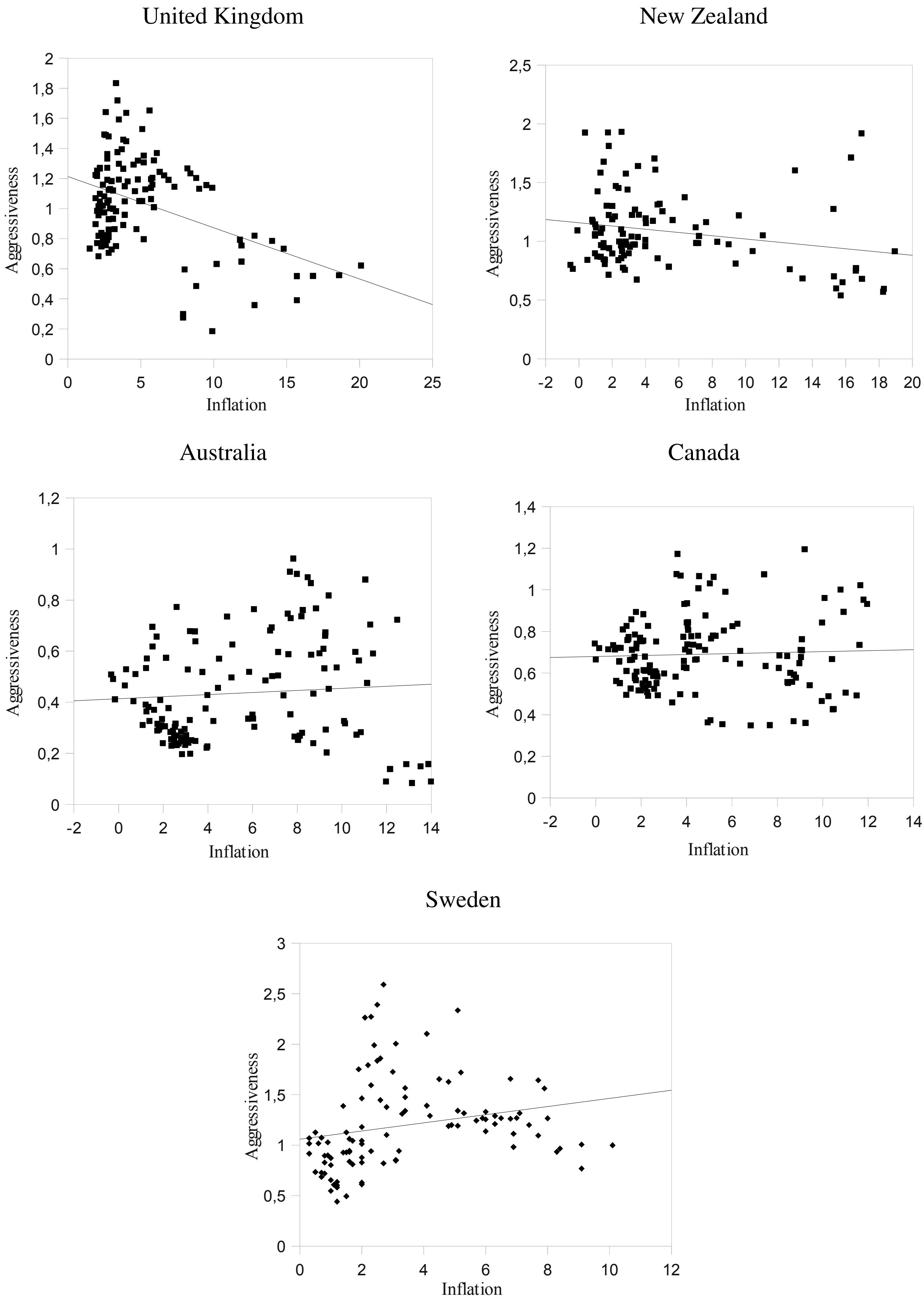
Figure 12 presents the monetary policy aggressiveness (defined as the estimate of the response of interest rates to expected inflation) in periods before and after the introduction of inflation targeting. It can be seen that in no country did the aggressiveness parameter increase after the adoption of inflation targeting. In fact, this aggressiveness substantially decreased in the UK, Australia and Sweden. If we look at the link between aggressiveness and expected inflation within the IT period, the results suggest that in most countries the aggressiveness is higher the more expected inflation deviates from its target. This broadly corresponds to the findings of Davradakis and Taylor (2006), who document a non-linear policy rule for the UK of a similar pattern. Similarly, Demers and Rodriguez (2002) in their analysis of monetary policy in Canada argue, that as long as inflation remains within the target band, it is possible to have a coefficient on inflation equal to zero as agents strongly believe that the monetary policy is credible. Nevertheless, insignificant responses to inflation are not common in the literature. With respect to the fact that inflation becomes more and more a forward-looking phenomenon, affected by expectations along with the interest rate setting, and considering stability of both inflation and interest rates, lower estimated response of interest rates to inflation becomes reasonable.

Figure 12 – Monetary Policy Aggressiveness and Inflation Targeting



Note: The y axis depicts the evolution of the estimated parameter β (the response of interest rates to expected inflation) and the x axis represents time, with the year of inflation targeting adoption denoted by a black vertical line. The values of β for New Zealand and Sweden are plotted on the right-hand axis.

Figure 13 – Monetary Policy Aggressiveness and Inflation Rate



Note: The figure presents the scatter plots between the response of interest rates to expected inflation (β), labeled as aggressiveness, and inflation rate at the horizon of monetary policy transmission (2 years).

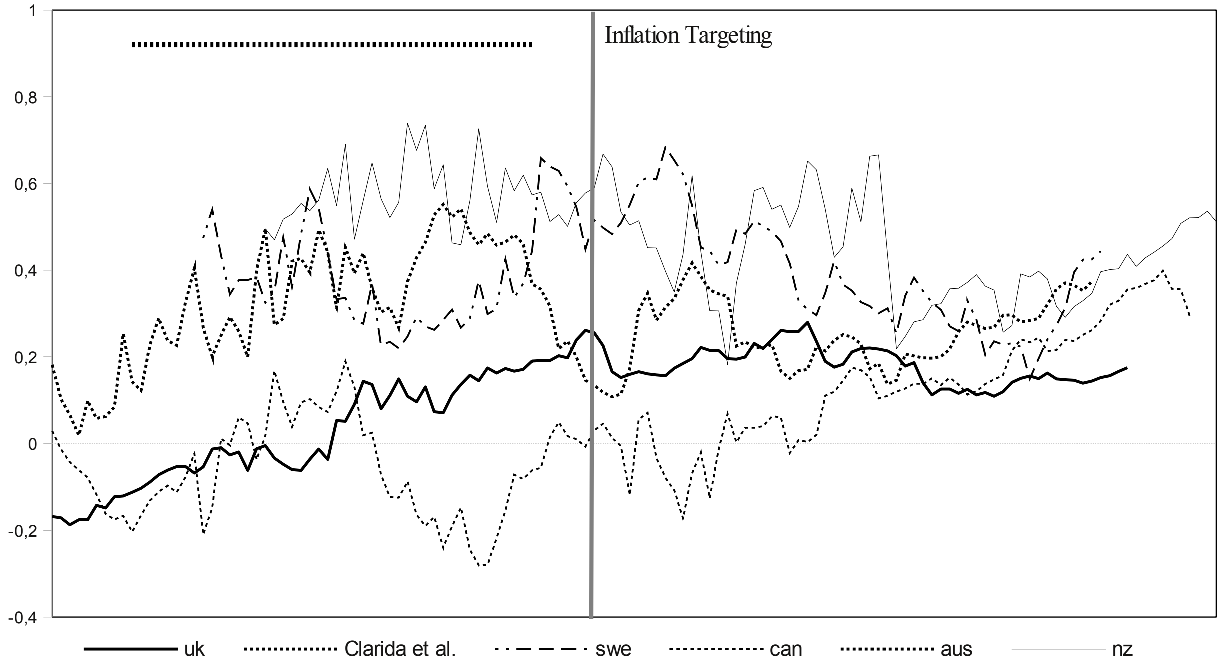
Figure 13 documents a link between the monetary policy aggressiveness and the inflation rate at the horizon of monetary policy transmission (two years ahead). As expected we find mainly negative relationship, the higher policy aggresiveness (coefficient β) determines lower inflation. The link is particulary strong in the UK and Canada. On the other hand, we find a counter-intuitive possitive relationship in Sweden.

The evolution of the estimated interest rate smoothing parameter in comparison to the time-invariant estimates for the UK in 1979–1990 by Clarida et al. (2000) is available in Figure 14. Our time-varying estimates of interest rate smoothing are well below the time-invariant one, which seems reasonable in the light of the recent critique by Rudebusch (2006), who puts forward that the degree of interest rate smoothing is actually low. While omitted variables or persistent shocks were deemed to be behind the implausible degree of policy inertia, our empirical results suggest that omission of the time-varying nature of the response coefficient may be another reason for the overestimation of smoothing coefficient ϱ in time-invariant policy rules. In time-varying framework, the time-varying intercept, that can be linked to a policy neutral rate, captures an important part of long-term dynamics of interest rates such as downward sloping trend in disinflation periods. Removing this information in fact plays a similar role as detrending usually does and persistence of deviations from time-varying trend decreases.

This hypothesis can be justified by observing results in the literature. Leigh (2008) estimates the FED implicit inflation target under the assumption of its time-varying nature. In terms of the Taylor rule it implies estimation of time varying intercept. Then, his smoothing parameter is 0,75 – lower than in most time-invariant studies. Boivin (2006) estimates the time-varying monetary policy rule for the United States and he reports lower values of inflation smoothing parameter for specifications.

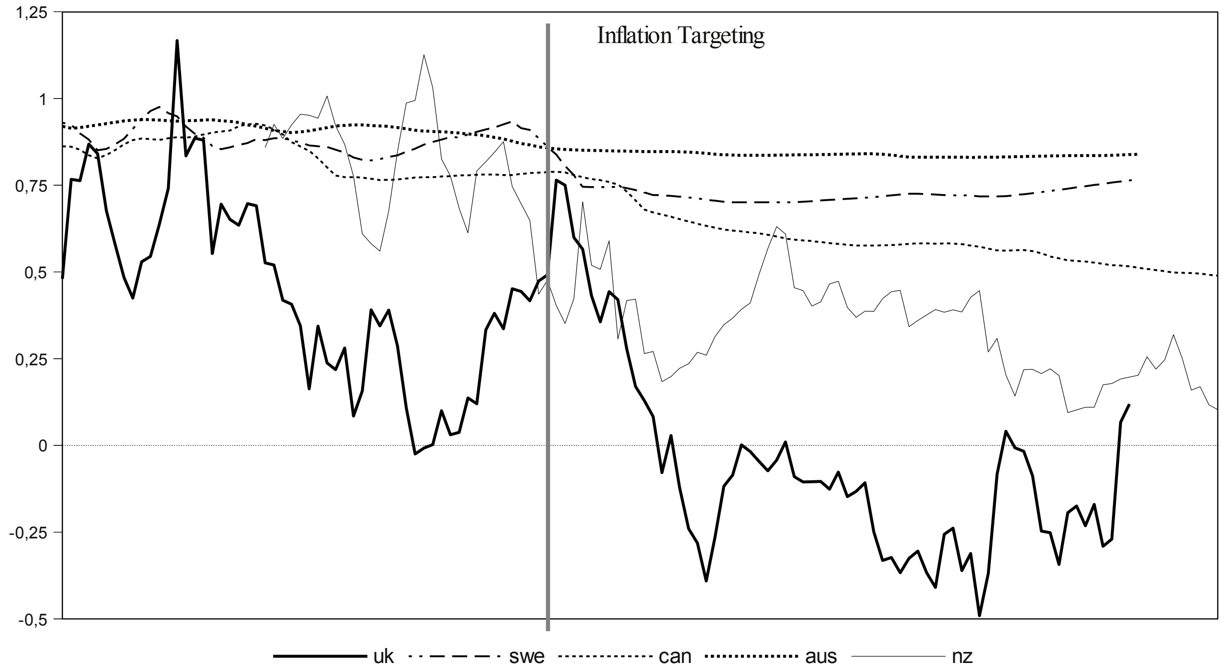
Finally, the results in Figure 15 plot the estimates of inflation persistence over time for all countries with respect to the inflation targeting adoption date. The results suggest that inflation persistence decreased after the adoption of inflation targeting, with a very distinct fall in the UK and New Zealand.

Figure 14 – Interest Rate Smoothing



Note: The figure presents the evolution of the estimated interest rate smoothing parameter ϱ over time in comparison to the interest rate smoothing parameter estimated in the time-invariant model of Clarida et al. (2000) for the UK.

Figure 15 – Inflation Targeting and Inflation Persistence



Note: The y axis depicts the evolution of the estimated inflation persistence parameter and the y axis represents time, with the year of inflation targeting adoption denoted by a black vertical line.

2.5 Conclusion

In this paper, we shed light on the evolution of monetary policy in the main inflation targeting central banks during the last three decades. The evolution of monetary policy is evaluated within a novel framework of a time-varying parameter model with endogenous regressors (Kim and Nelson, 2006), further addressing small sample issues (Schlicht, 1981; Schlicht, 2005; Schlicht and Ludsteck, 2006).

In our view, the results point to the usefulness of this econometric framework for analysis of the evolution of monetary policy setting. The estimation of standard monetary policy rules reveals that policy changes gradually and the changes coincide with several important institutional reforms as well as with the periods when the central banks successfully decreased double-digit inflation rates to rates consistent with their definitions of price stability.

In this respect, our results suggest that the response of interest rates to inflation is particularly high during periods when central bankers want to break a record of high inflation, such as in the UK in the early 1980s. Contrary to common perceptions, the response is often found to be less aggressive after the adoption of inflation targeting, suggesting a positive anchoring effect of this regime on inflation expectations. In other words, monetary policy need not be as aggressive as under a discretionary regime in order to achieve price stability (Kuttner and Posen, 1999). This result is supported by our finding that inflation becomes less inertial and the policy neutral rate arguably decreases after the adoption of inflation targeting.

We find that external factors matter for interest rate setting in all our sample countries. To be more precise, the foreign interest rate is found to enter the monetary policy rule significantly. The importance of the exchange rate varies, being apparently more important before the countries adopted inflation targeting than afterwards.

Our results also indicate that interest rate smoothing is much lower than typically reported by time-invariant estimates of monetary policy rules (see, for example, Clarida et al., 1998, 2000). Our estimates support the recent critique by Rudebusch (2006), who argues that the degree of interest rate smoothing is rather low. We suggest that neglect of changes in monetary policy setting over time is the reason for the implausible degree of policy inertia previously found.

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Appendix

Appendix 1 The Varying Coefficient Method

A.1.1 Introduction

In this section, we closely follow the Schlicht and Ludsteck (2006) paper. Consider a standard linear model:

$$y_t = a' x_t + u_t,\;\; a, x_t \in \mathbb{R}^n, u_t \sim N(0, \sigma^2),\;\; t = 1,2,\ldots T \qquad \text{(A.1)}$$

It can be extended for the case in which the coefficients a are allowed to follow a random walk. Then equation (A.1) is replaced by a system

$$y_t = a' x_t + u_t,\;\; u_t \sim N(0, \sigma^2) \qquad \text{(A.2)}$$

$$a_{t+1} = a_t + v_t,\;\; v_t \sim N(0, \Sigma) \qquad \text{(A.3)}$$

with one signal equation (A.2) and n state equations (A.3) for each time-varying parameter. The variance-covariance matrix Σ is assumed to be diagonal, that is

$$\Sigma = \begin{pmatrix} \sigma_1^2 & 0 & \ldots & 0 \\ 0 & \sigma_2^2 & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \ldots & \sigma_n^2 \end{pmatrix}$$

Define the following matrices:

$$X = \begin{pmatrix} x'_1 & 0 & \ldots & 0 \\ 0 & x'_2 & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \ldots & x'_n \end{pmatrix} \qquad P = \begin{pmatrix} -I_n & I_n & 0 & \ldots & 0 \\ 0 & -I_n & I_n & \ldots & 0 \\ \vdots & \vdots & \ddots & \ddots & \vdots \\ 0 & 0 & \ldots & -I_n & I_n \end{pmatrix}$$

of order T × Tn (T-1)n × Tn

$$y = \begin{pmatrix} y_1 \\ y_2 \\ \vdots \\ y_T \end{pmatrix},\quad u = \begin{pmatrix} u_1 \\ u_2 \\ \vdots \\ u_T \end{pmatrix},\quad a = \begin{pmatrix} a_1 \\ a_2 \\ \vdots \\ a_T \end{pmatrix},\quad v = \begin{pmatrix} v_2 \\ v_3 \\ \vdots \\ v_T \end{pmatrix}$$

of order T × 1 T ×1 Tn × 1 (T-1)n × 1

The system (A.2) and (A.3) can be rewritten as

$$y = Xa + u,\;\; u \sim N(0, \sigma^2 I_T) \qquad \text{(A.4)}$$

$$Pa = v,\;\; v \sim N(0, V), V = I_{T-1} \otimes \Sigma \qquad \text{(A.5)}$$

Estimation of the model based on equations (A.4) and (A.5) requires derivation of a distribution function that maps the random variables u_t and v_t to a set of observations X_t. However, such inference is not possible because the matrix P in (A.5) is of rank $(T-1)n$ rather than Tn and thus it cannot be inverted. Furthermore, any v does not determine the path of a_t uniquely.

A.1.2 The Orthogonal parametrization

The VC method used in this paper starts with an explicit definition of a set of possible values of a conditioned by matrix P and random variable v. Following the equation (A.5), any solution a can be written as

$$a = P'(PP')^{-1} v + Z\lambda \tag{A.6}$$

with $\lambda = \mathbb{R}^n$ and $Z = \frac{1}{\sqrt{T}} \begin{pmatrix} I_n \\ I_n \\ \vdots \\ I_n \end{pmatrix}$.

Therefore, the matrix Z is a matrix that translates the vector λ of dimension n to dimension Tn×1 (dimension of a) in order to get the same dimension of the first and the second term of the right hand side of the equation (A.6). The vector λ expresses additive shifts in the level of parameters a that leaves the disturbances v unaffected. Hence equation (A.5) becomes

$$w = XP'(PP')^{-1} v + u \tag{A.7}$$

Equations (A.6) and (A.7) build an orthogonal parametrization of the true model (A.4) and (A5). The orthogonally parametrized model implies that a_t follows a random walk and, that its path depends on all realizations of a random variable v_t.[29]

The equation (A.7) can be written as

$$y = XZ\lambda + w \tag{A.8}$$

where $w = XP'(PP')^{-1} v + u$.

The variable w is normally distributed:

$$w \sim N(0, W), \quad W = XBX' + \sigma^2 I_T \tag{A.9}$$

with $B = P'(PP')^{-1} V (PP')^{-1} P$.

Let the matrix of the observations follow a conventional format:

$$X^* = \sqrt{T}\, XZ = X \begin{pmatrix} I_n \\ I_n \\ \vdots \\ I_n \end{pmatrix} = \begin{pmatrix} x_1' \\ x_2' \\ \vdots \\ x_T' \end{pmatrix} \tag{A.10}$$

Inserting (A.12) into (A.8) implies a generalized linear regression model

$$y = \frac{1}{\sqrt{T}} X^* \lambda + w = X\beta + w \tag{A.11}$$

with $\hat{\beta} = \frac{1}{\sqrt{T}} \hat{\lambda}$.

The estimate of $\hat{\lambda}$ satisfies

$$\hat{\lambda} = (Z'X'W^{-1}XZ)^{-1} Z'X'W^{-1} y \tag{A.12}$$

29 To avoid excessive number o indexes, we skipped the time index t in the latter part of the text.

which is a standard GLS estimator of the classical regression problem with covariance matrix of residuals W and observations ZX. Taking expectations of a from A.6 and substituting $\hat{\lambda}$ for λ implies $Z'\hat{a}=\hat{\lambda}$ and hence $\frac{1}{T}\sum_{t=1}^{T} a_t=\beta$ in the GLS regression A.11.

A.1.3 Estimation of coefficients

The orthogonal parametrization derived in the previous section might be used for direct ML estimation of the time-varying parameters a. However, the derivation of the ML estimate of the vector of parameters a leads to a formulation that is equivalent to the minimization of the weighted sum of squares

$$\sum_{t=1}^{T} u^2+\theta_1\sum_{t=1}^{T} v_1^2+\theta_2\sum_{t=1}^{T} v_2^2+\ldots+\theta_n\sum_{t=1}^{T} v_n^2 \tag{A.13}$$

where the weights θ_i are the inverse variance ratios of the regression residuals u_t and the shocks in time-varying coefficients v_t, that is $\theta_i=\sigma^2/\sigma_i^2$. The proof can be found in Schlicht and Ludsteck, 2006, section 5. Hence the estimator balances the fit of the model and the parameter stability.[30]

Now we derive the formula used for estimation of the coefficients. For given X and y the estimated disturbances are

$$\begin{aligned}\hat{u}&=y-X\hat{a}\\ \hat{v}&=P\hat{a}\end{aligned} \tag{A.14}$$

Using the expressions for the estimated disturbances (14), minimization of the weighted sum of squares (13) implies

$$(X'X+\sigma^2P'V^{-1}P)\hat{a}=X'y \tag{A.15}$$

which is used for the estimation of coefficients $\hat{a}_t$ (Theorem 1, Schlicht-Ludsteck, 2006). The term in parentheses is a system matrix, M, of order $Tn\times Tn$. The estimator $\hat{a}$ is normally distributed with mean a and covariance $E\{(\hat{a}-a)^2\}=\sigma^2 M^{-1}$. The standard errors are then derived as the square roots of the main diagonal elements.

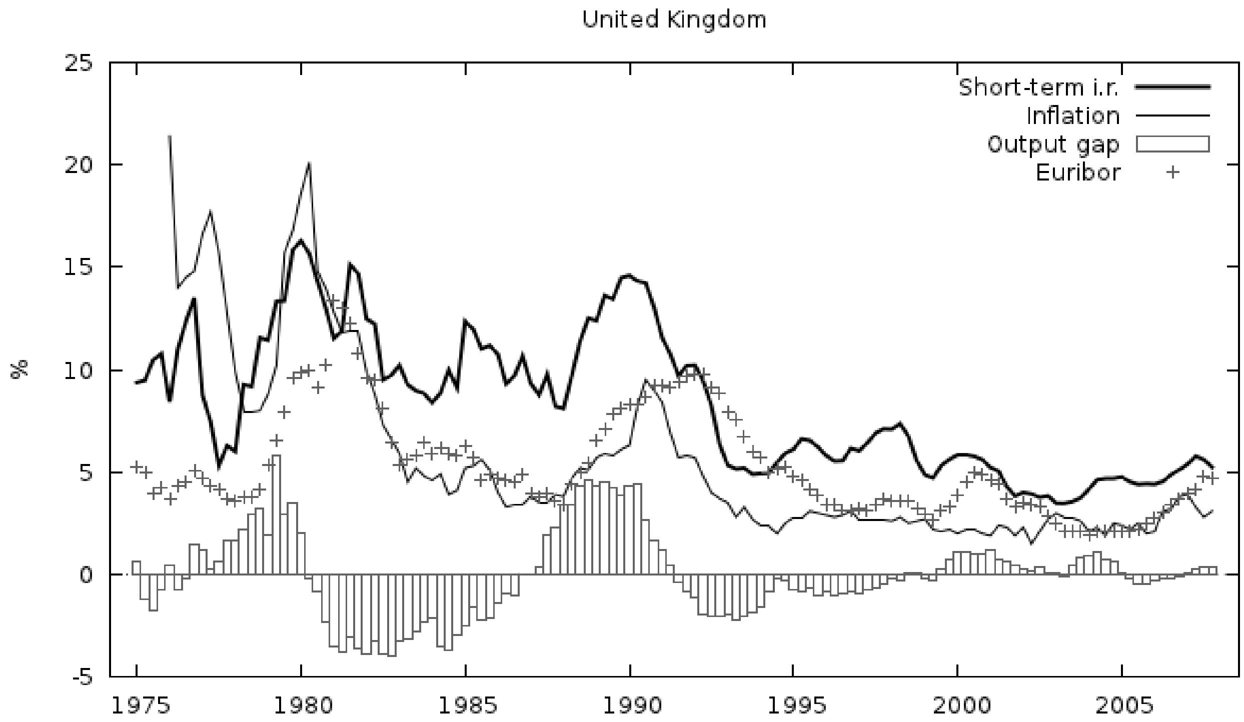
The coefficients estimated using the VC method have a straightforward interpretation: they have a time-invariant part, determined by a regression with fixed coefficients, and a random part reflecting the idea that some proportion of the variance of the dependent variable is caused by a change in the coefficients.

The estimation procedure proceeds as follows. The iterative procedure has two steps. First, given variances of residuals in both equations in (A.4) and (A.5), σ^2 and σ_i^2, the coefficients a_t are estimated using (A.15). Second, the estimated residuals are calculated using (A.14) and their estimated second moments $\hat{u}'\hat{u}$ and $\hat{v}'_i\hat{v}_i$ are compared to their expected moments $E[\hat{u}'\hat{u}]$ and $E[\hat{v}'_i\hat{v}_i]$. These steps are repeated until the estimated moments are identical to their expected counterparts (for a precise derivation of the moment estimator as well as computational details see Schlicht and Ludsteck, 2006, sections 6-9).

30 Originally, Schlicht and Ludsteck (2006) start with a derivation of the maximum likelihood estimator of parameters a based on the idea of orthogonal parameterization, which is described in the Appendix. Then they prove that the weighted least squares estimator is identical to the maximum likelihood estimator.

Appendix 2 Data

Figure A.1 Interest rates, Inflation and Output gap



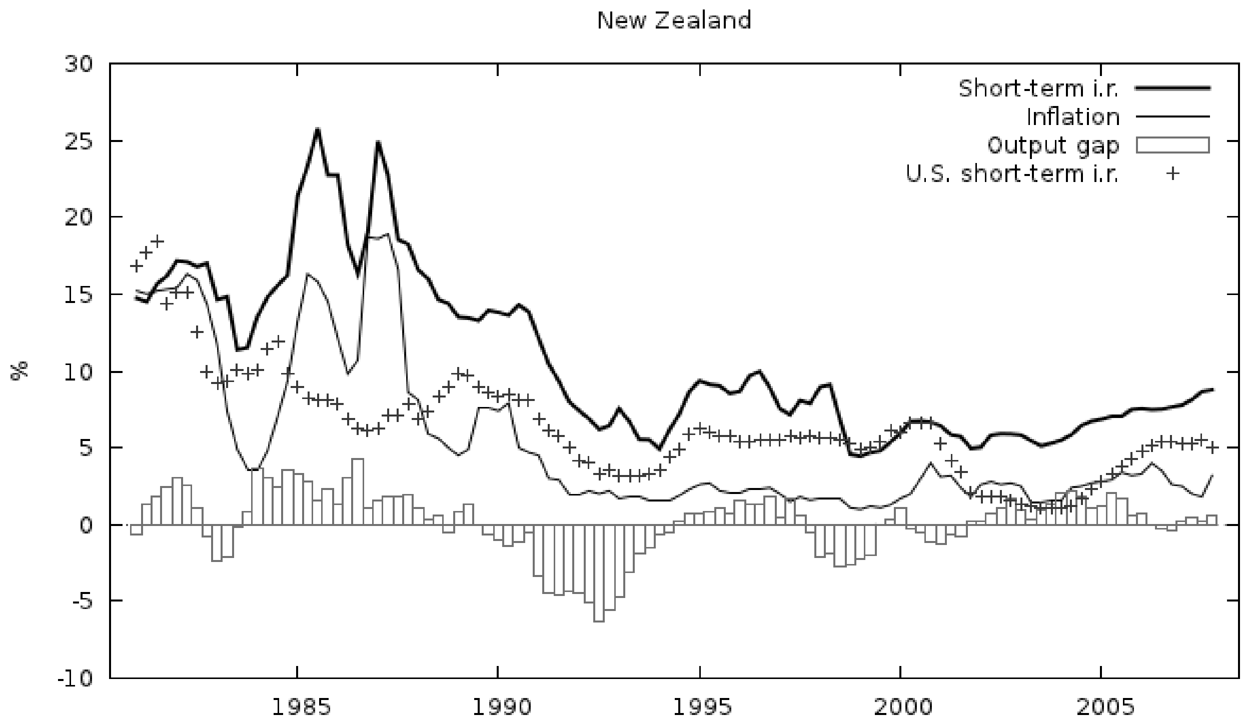
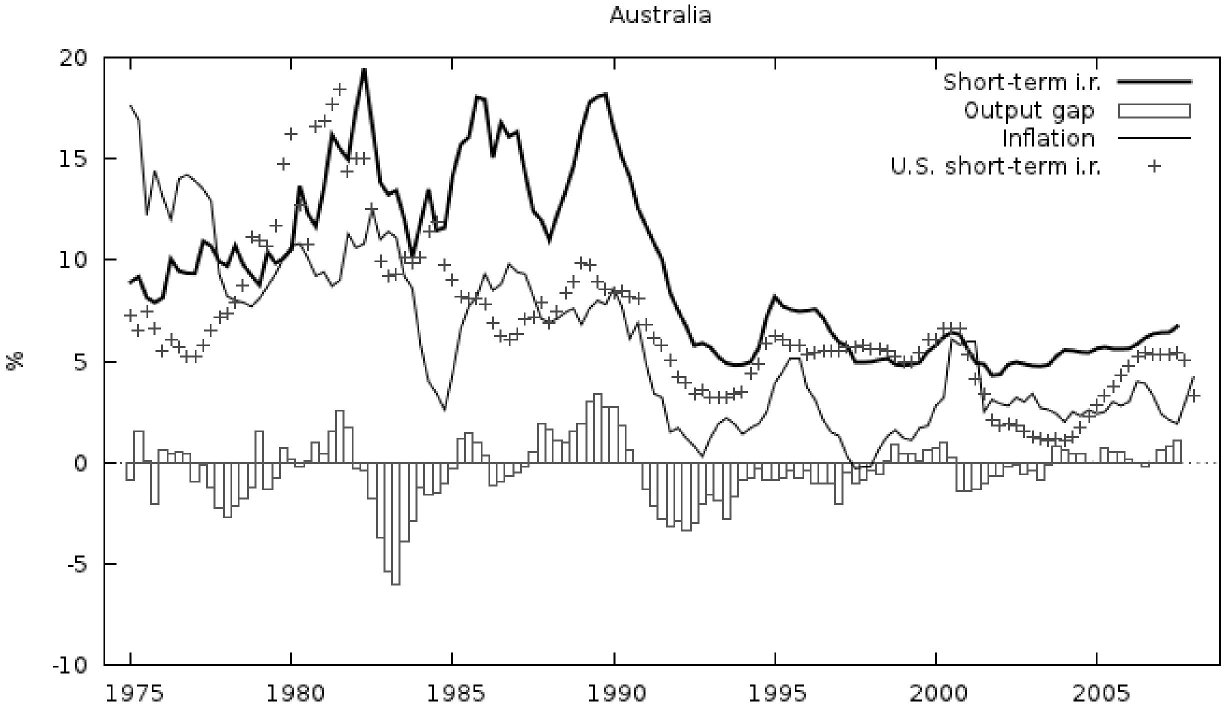


Figure A.1 Interest rates, Inflation and Output gap (Cont.)



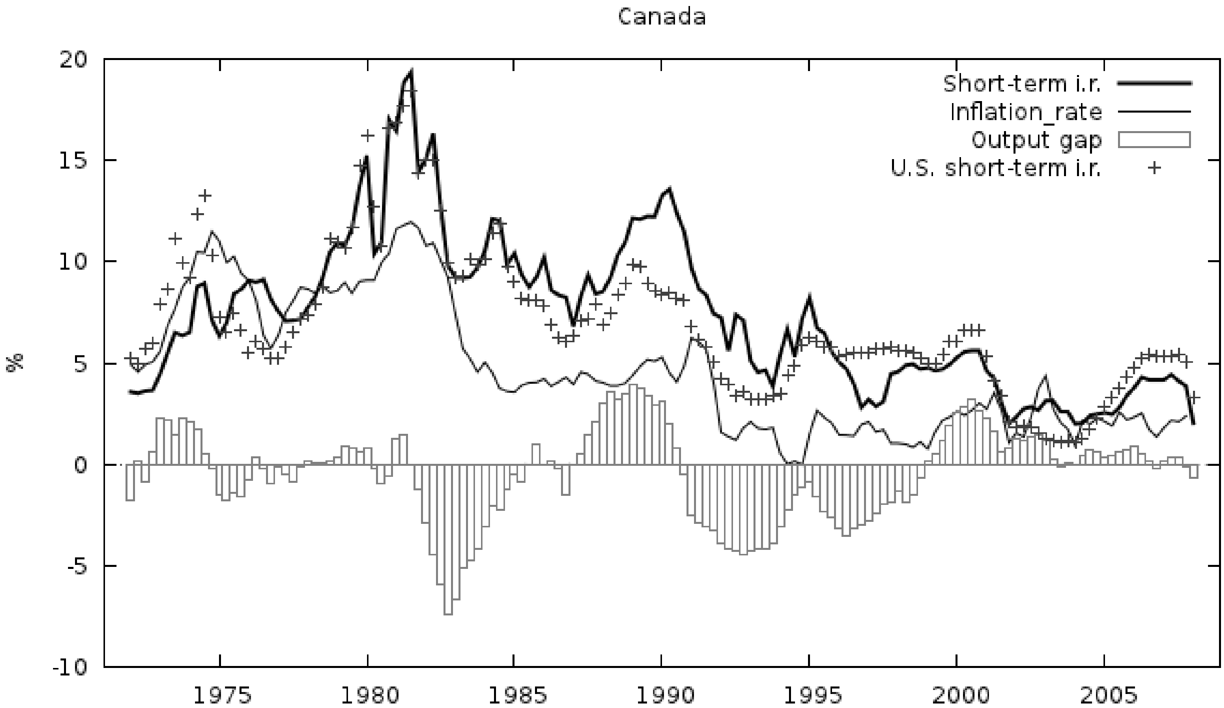


Figure A.1 Interest rates, Inflation and Output gap (Cont.)

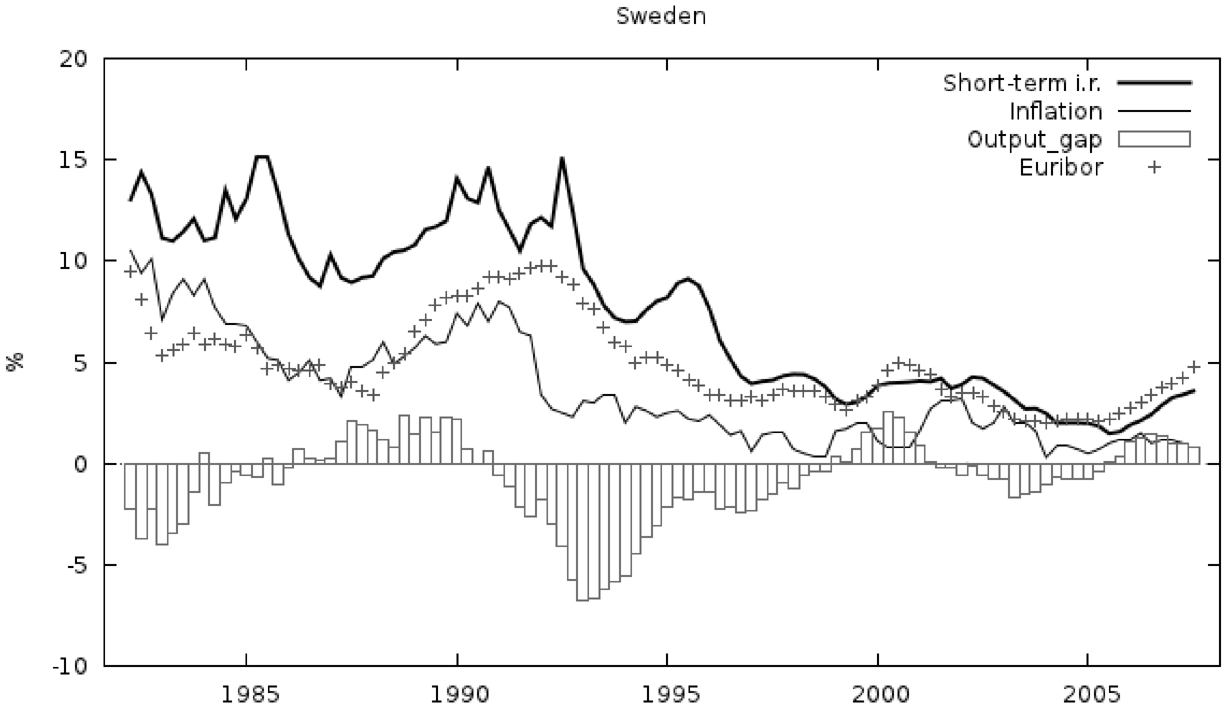
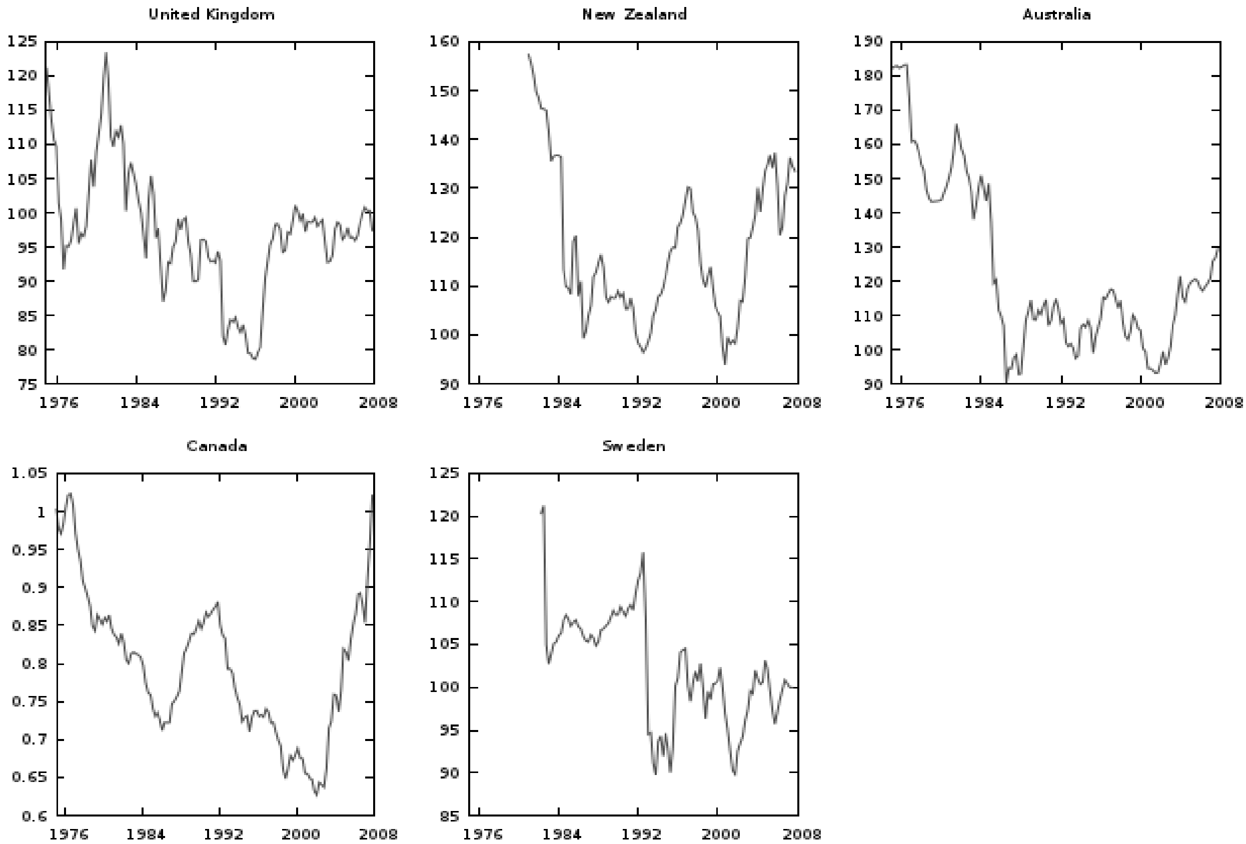
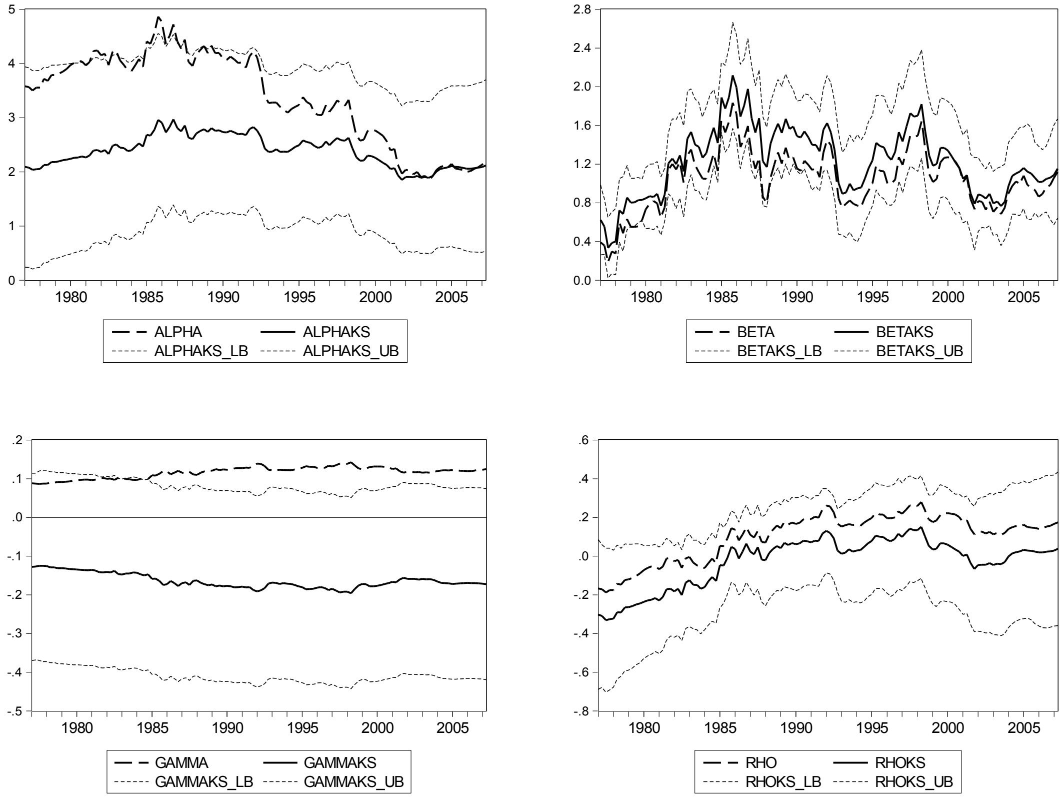


Figure A.2 Nominal effective exchange rates, CAD/USD exchange rate



Appendix 3 Additional Figures and Tables

Figure A.3 - Comparison of Estimated Coefficients by VC Method and the Kalman Smoother, UK



Note: 95% confidence bands for the Kalman smoother estimates. The upper-left graph depicts the time-varying intercept. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

Table A.1 Estimated Coefficients of Endogeneity Correction Terms

		UK	NZ	AUS	CAN	SWE
Inflation	mean	**-0.823**	-0.396	-0.160	**-1.142**	**-0.557**
	s.e.	0.320	0.403	0.180	0.433	0.268
Output gap	mean	0.239	-0.040	0.030	**-0.528**	-0.027
	s.e.	0.240	0.153	0.189	0.127	0.134

Note: Bold: sign. at 5%

Chapter 3

Time-Varying Monetary-Policy Rules and Financial Stress: Does Financial Instability Matter for Monetary Policy?

3.1 Introduction

The recent financial crisis has intensified the interest in exploring the interactions between monetary policy and financial stability. Official interest rates were driven sharply to historical lows, and many unconventional measures were used to pump liquidity into the international financial system. Central banks pursued monetary policy under high economic uncertainty coupled with large financial shocks in many countries. The financial crisis also raised new challenges for central bank policies, in particular the operationalization of issues related to financial stability for monetary-policy decision making (Goodhart, 2006; Borio and Drehmann, 2009).

This paper seeks to analyze whether and how monetary policy interest rates evolved in response to financial instability over the last three decades. The monetary policies of central banks are likely to react to financial instability in a non-linear way (Goodhart et al., 2009). When a financial system is stable, the interest-rate-setting process largely reflects macroeconomic conditions, and financial stability considerations enter monetary policy discussions only to a limited degree. On the other hand, central banks may alter their monetary policies to reduce financial imbalances if these become severe. In this respect, Mishkin (2009) questions the traditional linear-quadratic framework[1] when financial markets are disrupted and puts forward an argument for replacing it with non-linear dynamics describing the economy and a non-quadratic objective function resulting in non-linear optimal policy.

To address the complexity of the nexus between monetary policy and financial stability as well as to evaluate monetary policy in a systematic manner, this paper employs the recently developed time-varying parameter estimation of monetary-policy rules, appropriately accounting for endogeneity in policy rules. This flexible framework, together with a new comprehensive financial stress dataset developed by the International Monetary Fund, will allow not only testing of whether central banks responded to financial stress, but also quantification of the magnitude of this response and detection of the periods and types of stress that were the most worrying for monetary authorities.

Although theoretical studies disagree about the role of financial instability for central banks' interest-rate-setting policies, our empirical estimates of the time-varying monetary-policy rules of the US Fed, the Bank of England (BoE), the Reserve Bank of Australia (RBA), the Bank of Canada (BoC), and Sveriges Riksbank (SR) show that central banks often alter the course of monetary policy in the face of high financial stress, mainly by decreasing policy rates.[2] However, the size of

1 That is, linear behavior of the economy and a quadratic objective function of the monetary authority.

2 Our choice of countries is based on data availability and on the suitability of the data for our econometric framework. Due to limited data availability, we do not include the Reserve Bank of New Zealand, the ECB, and emerging countries. The Bank of Japan could not be included either, given that its policy rates were flat for an extended period.

this response varies substantially over time as well as across countries. There is some cross-country and time heterogeneity as well when we examine central banks' considerations of specific types of financial stress: most of them seemed to respond to stock-market stress and bank stress, and exchange-rate stress drives central bank reactions only in more open economies

The paper is organized as follows. Section 2 discusses related literature. Section 3 describes our data and empirical methodology. Section 4 presents our results. Section 5 concludes. An appendix with a detailed description of the methodology and additional results follows.

3.2 Related Literature

First, this section gives a brief overview of the theory as well as empirical evidence on the relationship between monetary policy (rules) and financial instability. Second, it provides a short summary of various measures of financial stress.

3.2.1 Monetary policy (rules) and financial instability – some theories

Financial friction, such as unequal access to credit or debt collateralization, is recognized as having important consequences for monetary policy transmission, and Fisher (1933) has already presented the idea that adverse credit-market conditions can cause significant macroeconomic disequilibria.

During the last two decades, the effects of monetary policy have been studied mainly within New Keynesian (NK) dynamic stochastic general equilibrium (DSGE) models, which assume the existence of nominal rigidities. The common approach to incorporating financial market friction within the DSGE framework is to introduce the financial accelerator mechanism (Bernanke et al., 1996, 1999), implying that endogenous developments in credit markets work to amplify and propagate shocks to the macro economy. Tovar (2009) emphasizes that the major weakness of the financial accelerator mechanism is that it only addresses one of many possible financial frictions. Goodhart et al. (2009) note that many NK DSGE models lack the financial sector completely or model it in a rather embryonic way. Consequently, more recent contributions within this stream of literature have examined other aspects of financial friction, such as balance sheets in the banking sector (Choi and Cook, 2004), the portfolio-choice issue with complete (Engel and Matsumoto, 2009) or incomplete markets (Devereux and Sutherland, 2007), and collateral constraints (Iacovello and Neri, 2010).[3]

A few studies focus more specifically on the relationship between the monetary-policy stance (or the monetary-policy rule) and financial stability. However, they do not arrive at a unanimous view of whether a monetary-policy rule should include some measure of financial stability. Brousseau and Detken (2001) present an NK model where a conflict arises between short-term price stability and financial stability due to a self-fulfilling belief linking the stability of inflation to the smoothness of the interest-rate path and suggests that monetary policy should react to financial instability. Akram et al. (2007) investigate the macroeconomic implications of pursuing financial stability within a flexible inflation-targeting framework. Their model, using a policy rule augmented by financial-stability indicators, shows that the gains of such an augmented rule vis-à-vis the rule without financial-stability indicators highly depends on the nature of the shocks. Akram and Eitrheim (2009) build on the previous framework, finding some evidence that the policy response to housing prices, equity prices or credit growth can cause high interest-rate volatility and

3 A survey of this literature is provided by Tovar (2009).

actually lower financial stability in terms of indicators that are sensitive to interest rates. Cecchetti and Li (2008) show, in both a static and dynamic setting, that a potential conflict between monetary policy and financial supervision can be avoided if the interest-rate rule takes into account (procyclical) capital-adequacy requirements, in particular, that policy interest rates are lowered when financial stress is high. Bauducco et al. (2008) extend the current benchmark NK model to include financial systems and firms that require external financing. Their simulations show that if a central bank responds to financial instability by policy easing, it achieves better inflation and output stabilization in the short term at the cost of greater inflation and output volatility in the long term, and vice versa. For the US Fed, Taylor (2008) proposes a modification of the standard Taylor rule to incorporate adjustments to credit spreads. Teranishi (2009) derives a Taylor rule augmented by the response to credit spreads as an optimal policy under heterogeneous loan-interest-rate contracts. He finds that the policy response to a credit spread can be both positive and negative, depending on the financial structure. However, he also proposes that when nominal policy rates are close to zero, a commitment rather than a discretional policy response is the key to reducing credit spreads. Christiano et al. (2008) suggest augmenting the Taylor rule with aggregate private credit and find that such a policy would raise welfare by reducing the magnitude of the output fluctuations. Cúrdia and Woodford (2010) develop a NK DSGE model with credit friction to evaluate the performance of alternative policy rules that are augmented by a response 1) to credit spreads and 2) to aggregate the volume of private credit in the face of different shocks. They argue that the response to credit spreads can be welfare improving, but the optimal size of such a response is probably rather small. Like Teranishi (2009), they find little support for augmenting the Taylor rule by the credit volume, given that the size and even the sign of the desired response is sensitive to the sources of shock and their persistence, which is information that is not always available during operational policy making.

A related stream of literature focuses on the somewhat narrower issue of whether or not monetary policy should respond to asset prices. Bernanke and Gertler (1999, 2001) argue that the stabilization of inflation and output provides a substantial contribution to financial stability and that there are few, if any, gains to responding to asset prices. Faia and Monacelli (2007) extend the model developed by Bernanke and Gertler (2001) by a robust welfare metric, confirming that strict inflation stabilization offers the best solution. Cecchetti et al. (2000) take the opposite stance, arguing that developments in asset markets can have a significant impact on both inflation and real economic activity, and central banks might achieve better outcomes by considering asset prices provided they are able to detect asset-price misalignments. Borio and Lowe (2002) support this view, claiming that financial imbalances can build up even in a low-inflation environment, which is normally favorable to financial stability. The side effect of low inflation is that excess demand pressures may first appear in credit aggregates and asset prices rather than consumer prices, which are normally considered by policy makers. Gruen et al. (2005) argue that responding to an asset bubble is feasible only when the monetary authority is able to make a correct judgment about the process driving the bubble. Roubini (2006) and Posen (2006) provide a summary of this debate from a policy perspective.

3.2.2 Monetary policy (rules) and financial instability – empirical evidence

The empirical evidence on central banks' reactions to financial instability is rather scant. Following the ongoing debate about whether central banks should respond to asset-price volatility (e.g. Bernanke and Gertler, 1999, 2001; Cecchetti et al., 2000; Bordo and Jeanne, 2002), some studies have tested the response of monetary policy to different asset prices, most commonly stock prices (Rigobon and Sack, 2003; Chadha et al., 2004; Siklos and Bohl, 2008; Fuhrer and Tootell, 2008). They find some evidence either that asset prices entered the policy-information set (because they contain information about future inflation) or that some central banks were directly trying to offset these disequilibria.[4] All of these papers estimate time-invariant policy rules, which means that they test a permanent response to these variables. However, it seems more plausible that if central banks respond to asset prices, they do so only when asset-price misalignments are substantial; in other words, their responses are asymmetric. There are two additional controversies related to the effects of asset prices on monetary-policy decisions. The first concerns the measure, in particular whether the stock-market index that is typically employed is sufficiently representative, or whether some other assets, in particular housing prices, should be considered as well. The second issue is related to the (even ex-post) identification of asset-price misalignment. Finally, it is likely that the perception of misalignments is influenced by general economic conditions and that a possible response might evolve over time.

Detken and Smets (2004) summarize some stylized facts on macroeconomic and monetary-policy developments during asset-price booms. Overall, they find that monetary policy was significantly looser during high-cost booms that were marked by crashes of investment and real-estate prices in the post-boom periods.

A few empirical studies measure the monetary-policy response using broader measures of financial imbalances. Borio and Lowe (2004) estimate the response of four central banks (the Reserve Bank of Australia, the Bundesbank, the Bank of Japan, and the US Fed) to imbalances proxied by the ratio of private-sector credit to GDP, inflation-adjusted equity prices, and their composite. They find either negative or ambiguous evidence for all countries except the USA, confirming that the Fed responded to financial imbalances in an asymmetric and reactive way, i.e., that the federal funds rate was disproportionately lowered in the face of imbalance unwinding, but was not tightened beyond normal as imbalances built up. Cecchetti and Li (2008) estimate a Taylor rule augmented by a measure of banking stress, in particular the deviation of leverage ratios (total loans to the sum of equity and subordinated debt; total assets to the sum of bank capital and reserves) from their Hodrick-Prescott trend. They find some evidence that the Fed adjusted the interest rate to counteract the procyclical impact of a bank's capital requirements, while the Bundesbank and the Bank of Japan did not. Bulíř and Čihák (2008) estimate the monetary-policy response to seven alternative measures of financial-sector vulnerability (crisis probability, time to crisis, distance to default or credit default swap spreads) in a panel of 28 countries. Their empirical framework is different in the sense that the monetary-policy stance is proxied along the short-term interest rate by measures of domestic liquidity, and external shocks are controlled for. In the panel setting, they find a statistically significant negative response to many variables representing vulnerability (policy easing) but, surprisingly, not in country-level regressions. Belke and Klose (2010) investigate the factors behind the interest-rate decisions of the ECB and the Fed during the current crisis. They conclude that the estimated policy rule was significantly altered only for the Fed, and they put forward that the ECB gave greater weight to inflation stabilization at the cost of some output loss.

4 A similar but somewhat less polemic debate applies to the role of exchange rates, especially for small, open economies (Taylor, 2001).

3.2.3 Measures of financial stress

The incidence and determinants of different types of crises have been typically traced in the literature by a means of narrative evidence (expert judgment). This has sometimes been complemented by selected indicators (exchange rate devaluation or the state of foreign reserves) that point to historical regularities (e.g., Eichengreen and Bordo, 2002; Kaminsky and Reinhart, 1999; Reinhart and Rogoff, 2008; Laeven and Valencia, 2008). The empirical studies (e.g., Goldstein et al., 2000) used binary variables that were constructed based on these narratives.

Consequently, some contributions strived to provide more data-driven measures of financial stress. Most of the existing stress indices are based on high-frequency data, but they differ in the selected variables (bank capitalization, credit ratings, credit growth, interest rate spreads or volatility of different asset classes), country coverage, and the aggregation method. An important advantage of continuous stress indicators is that they may reveal periods of small-scale stress that did not result in full-blown crises and were neglected in studies based on binary crisis variables.

The Bank Credit Analyst (BCA) reports a monthly financial stress index (FSI) for the USA that is based on the performance of banking shares compared to the whole stock market, credit spreads and the slope of the yield curve, and new issues of stocks and bonds and consumer confidence. JP Morgan calculates a Liquidity, Credit and Volatility Index (LCVI) based on seven variables: the US Treasury curve error (the standard deviation of the spread between on-the-run and off-the-run US Treasury bills and bonds along the entire maturity curve), the 10-year US swap spread, US high-yield spreads, JP Morgan's Emerging Markets Bond Index, foreign exchange volatility (the weighted average of the 12-month implied volatilities of several currencies), the Chicago Board of Exchange VIX equity volatility index, and the JP Morgan Global Risk Appetite Index.

Illing and Liu (2006) develop a comprehensive FSI for Canada. Their underlying data cover equity, bond, and foreign exchange markets as well as the banking sector. They use a standard measure and refined measure of each stress component, where the former refers to the variables and their transformations that are commonly found in the literature, while the latter incorporates adjustments that allow for better extraction of information about stressful periods. They explore different weighting schemes to aggregate the individual series (factor analysis, the size of the corresponding market for total credit in the economy, variance-equal weighting). Finally, they perform an expert survey to identify periods that were perceived as especially stressful, confirming that the FSI matches these episodes very well.

For the Fed Board of Governors, Carlson et al. (2008) propose a framework similar to the option-pricing model (Merton, 1974) that aims to provide the distance-to-default of the financial system, the so-called Index of Financial Health. The method uses the difference between the market value of a firm's assets and liabilities and the volatility of the asset's value to measure the proximity of a firm's assets to being exceeded by their liabilities. They apply this measure to 25 of the largest US financial institutions, confirming its impact on capital investments in the US economy. The Kansas City Fed developed the Kansas City Financial Stress Index (Hakkio and Keeton, 2009), which is published monthly and is based on eleven variables (seven spreads between different bond classes by issuers, risk profiles and maturities, correlations between returns on stocks and Treasury bonds, expected volatility of overall stock prices, volatility of bank stock prices, and a cross-section dispersion of bank stock returns) that are aggregated by principal component analysis.

Finally, the International Monetary Fund (IMF) recently published financial stress indices for various countries. Cardarelli et al. (2011) propose a comprehensive index based on high-frequency data where the price changes are measured with respect to their previous levels or trend values. The underlying variables are standardized and aggregated into a single index (FSI) using variance-equal weighting for each country and period. The FSI has three subcomponents: the banking sector (the slope of the yield curve, TED spread, and the beta of banking-sector stocks), securities markets (corporate bond spreads, stock-market returns and time-varying volatility of stock returns) and exchange rates (time-varying volatility of NEER changes). Balakrishnan et al. (2009) modify the previous index to account for the specific conditions of emerging economies, on the one hand including a measure of exchange rate pressures (currency depreciation and decline in foreign reserves) and sovereign debt spread, and on the other hand downplaying the banking-sector measures (slope of the yield curve and TED spread).[5] We will use the former index, given its comprehensiveness as well as its availability for different countries (see more details below).

3.3 Data and Empirical Methodology

3.3.1 The dataset

Given the frequency of monetary policy committee meetings in most central banks, we use monthly data (due to unavailability of all monthly series for a sufficiently long time period, we use quarterly data for Sweden and Canada). The sample periods vary slightly due to data availability (the US 1981:1M–2009:6M; the UK 1981:1M–2009:3M; Australia 1983:3M–2009:5M; Canada 1981:1Q–2008:4Q; Sweden 1984:2Q–2009:1Q).

The dependent variable is typically an interest rate closely related to the official (censored) policy rate, in particular the federal funds rate (3M) for the USA, the discount rate (three-month Treasury bills) for the UK, Canada, and Sweden, and the three-month RBA-accepted bills rate for Australia. It is evident that the policy rate is not necessarily the only instrument that central banks use, especially during the 2008–2009 global financial crisis, when many unconventional measures were implemented (see Borio and Disyatat, 2009; Reis, 2010). To address this issue in terms of estimated policy rules, for a robustness check we use the interbank interest rate (at a maturity of three months). While both rates are used in empirical papers on monetary-policy rule estimation without great controversy, the selection of the interest rate becomes a more delicate issue during periods of financial stress (Taylor, 2008). While the former is more directly affected by genuine monetary-policy decisions (carried out by open market operations), the latter additionally includes liquidity conditions on interbank markets and, as such, can be affected by unconventional policies, though these are usually insulated (often intentionally) from policy interest rates.[6] This is a drawback but also a potential advantage of this alternative dependent variable. On the one hand,

5 The IMF FSI has recently been applied by Melvin and Taylor (2009) to analyze exchange rate crises.

6 Borio and Disyatat (2009) characterize unconventional policies as policies that affect the central bank's balance sheet size and composition and that can be insulated from interest rate policy (the so-called "decoupling principle"). One common example of such a policy (not necessarily used during times of crisis) is sterilized exchange-rate intervention. Given that we are looking not at a single episode of stress, but rather want to identify whether monetary authorities deviated from systematic patterns (the policy rule) during these periods (by responding to indicators of financial stress), we need to use a consistent measure of policy action that is adjusted during periods of financial stress, though other measures may be in place as well. Therefore, we assume that the monetary-policy stance is fully reflected in the interest rate, and we are aware that it might be subject to downward bias on the financial-stress coefficient. The reader may want to interpret our results on the importance of financial stress for interest-rate setting as a conservative estimate.

changes in official policy rates may not pass through fully to interbank interest rates, in particular when the perceived counterparty risk is too high and credit spreads widen (see Taylor and Williams, 2009). On the other hand, the interbank rate may also incorporate the impact of policy actions, such as quantitative easing aimed at supplying additional liquidity into the system.[7]

Inflation is measured as the year-on-year change in the CPI, apart from for the United States, where we use the personal consumption expenditures price index (PCE), and Sweden, where underlying CPIX inflation (which excludes households' mortgage-interest expenditures and the direct effects of changes in indirect taxes and subsidies from the CPI) is used.[8] The output gap is proxied by the gap of the seasonally adjusted industrial production index derived by the Hodrick-Prescott filter with a smoothing parameter set to 14,400.[9] For Sweden and Canada, where we use quarterly data, the output gap was taken as reported in the OECD Economic Outlook (production function method based on NAWRU — non-accelerating wage rate of unemployment).

We proxy financial stress by means of the FSI provided recently by the IMF (Cardarelli et al., 2011), which is a consistent measure for a wide range of countries but, at the same time, is sufficiently comprehensive to track stress of a different nature. It includes the main components of financial stress in an economy and is available for a reasonably long period to be used for our empirical analysis (see Figure 1). We use both the overall index, which is a sum of seven components, as well as each sub-index and component separately:

i. Banking-related sub-index components: the inverted term spread (the difference between short-term and long-term government bonds), TED spread (the difference between interbank rates and the yield on Treasury bills), banking beta (12-month rolling beta, which is a measure of the correlation of banking stock returns to total returns in line with the CAPM);

ii. Securities-market-related sub-index components: corporate bond spread (the difference between corporate bonds and long-term government bond yields), stock-market returns (monthly returns multiplied by -1), time-varying stock-return volatility from the GARCH(1,1) model;

iii. Foreign-exchange-related sub-index: the time-varying volatility of monthly changes in NEER, from the GARCH (1,1) model.

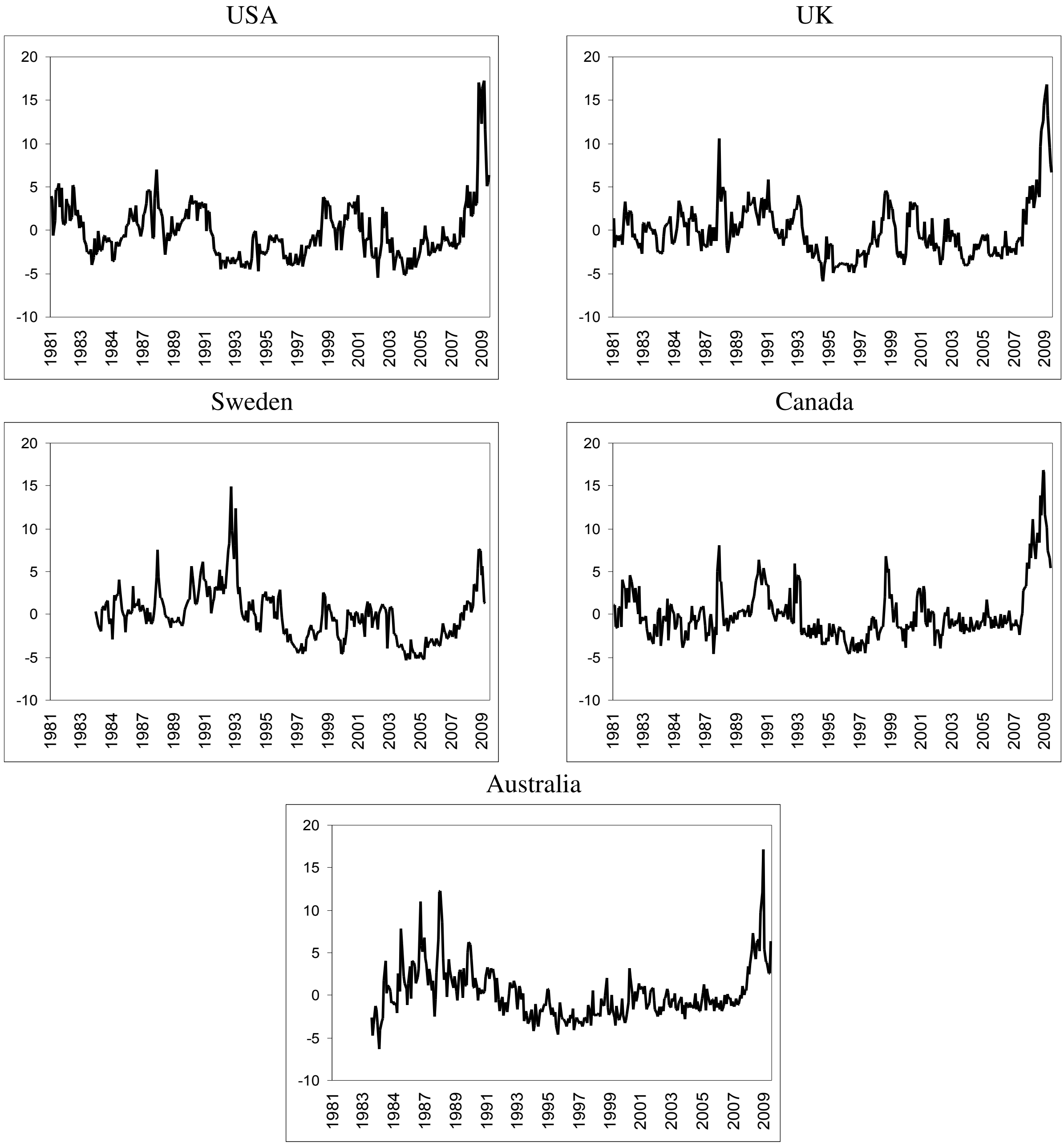
We examined various alternative methods of aggregating the components – simple sum, variance-equal weighting, and PCA weighting – but failed to uncover any systematic differences among these in terms of the values of the overall index and consecutively in the empirical results. Cardarelli et al. (2011) confirm that extreme values of this indicator correctly identify almost all (approximately 80%–90%) of the financial crises (including banking, currency, and other crises, along with stock and house-price boom and busts) identified in previous studies.

7 There are other policy measures that can be used as a reactive or pre-emptive response to financial stress, such as regulatory or administrative measures, although their effects are likely to appear only in the longer term and cannot be reasonably included in our empirical analysis.

8 For Australia, the monthly CPI is not available because both the Reserve Bank of Australia and the Australian Bureau of Statistics only publish quarterly data. The monthly series was obtained using linear interpolation of the CPI index.

9 The industrial production cycle had to be used as a proxy for the output gap given that GDP data are not available at monthly frequency. Though a bit more volatile, it is highly correlated with the output gap from GDP (comparison at quarterly frequency). Moreover, industrial production data tend to be revised less often and to a lesser extent than the GDP data, which reduced the problem of real-time vs. ex-post data present in the GDP data.

Figure 1 – IMF Financial Stress Indicator



Note: The figure presents the evolution of the IMF stress index over time.
Higher numbers indicate higher stress (see Cardarelli et al., 2011).

The use of a composite index has a number of benefits. First, it approximates the evolution of financial stress caused by different factors and thus is not limited to one specific type of instability. Second, the inclusion of additional variables in the stress index does not affect the evolution of the indicator markedly (Cardarelli et al., 2011). Third, the composition of the indicator allows for breaking down the reactions of the central bank with respect to different stress subcomponents. Nevertheless, one has to be cautious about the interpretation. The composite indicator might suggest a misleading interpretation as long as the stress is caused by variables not included in the FSI but rather highly correlated with some subcomponent. An example is the case of Sweden during the ERM crisis. At the time of the crisis, Sweden maintained a fixed exchange rate, and the Riksbank sharply increased interest rates to sustain the parity. However, this is not captured by the exchange-rate subcomponent of the FSI, which measures exchange-rate volatility, because the volatility was actually close to zero. A closer examination of the data shows that this period of stress is captured by the inverted term structure; hence, it is incorrectly attributed to bank stress. A similar pattern can be observed for the UK, where the FSI increases after the announcement of withdrawal from the ERM.

3.3.2 The empirical model

Following Clarida et al. (1998, 2000), most empirical studies assume that the central bank sets the nominal interest rate in line with the state of the economy typically in a forward-looking manner:

$$r_t^* = \bar{r} + \beta\left(E[\pi_{t+i}|\Omega_t] - \pi_{t+i}^*\right) + \gamma E[y_t|\Omega_t] \tag{1}$$

where r_t^* denotes the targeted interest rate, $\bar{r}$ is the policy neutral rate[10], π_{t+i} stands for the central bank forecast of the yearly inflation rate, i indicates periods ahead based on an information set Ω_t used for interest-rate decisions available at time t, and is π_{t+i}^* the central bank's inflation target.[11] y_t represents a measure of the output gap.

Nevertheless, Eq. (1) was found to be too restrictive to provide a reasonable description of actual interest-rate setting. Notably, it does not account for interest-rate smoothing by central banks, in particular the practice whereby the central bank adjusts the interest rate sluggishly to the targeted value. This is tracked in empirical studies by the simple partial-adjustment mechanism:

$$r_t = \rho r_{t-1} + (1-\rho) r_t^* \tag{2}$$

where $\rho \in [0,1]$ is the smoothing parameter. There is an ongoing controversy as to whether this parameter represents genuine policy inertia or reflects empirical problems related to omitted variables, dynamics or shocks (see, e.g., Rudebusch, 2006). The linear policy rule in Eq. (1) can be obtained as the optimal monetary-policy rule in the LQ framework, where the central bank aims only at price stability and economic activity. Bauducco et al. (2008) propose an NK model with a financial system where the central bank has privileged information (given its supervisory function) on the health of the financial sector. In such a setting, the common policy rule represented by Eq. (1) will be augmented by variables representing the health of the financial sector. Following this contribution, we consider the forward-looking rule where central banks may respond to a comprehensive measure of financial stress rather than stress in a particular segment (Bulíř and Čihák, 2008). In practice, the augmented rule can be of some interest to outsiders because inflation expected by the individual monetary-policy committee members is unobservable to the public (even though some central banks publish figures that may be very close to the unobserved expected inflation, such as staff inflation forecasts or inflation forecasts stemming from interactions between staff and monetary-policy committee members). In such case, outsiders may benefit from including additional indicators such as financial stress in the policy rule to predict the central bank's behavior more accurately.

Therefore, we substitute Eq. (2) into Eq. (1), eliminate unobserved forecast variables and include measures of the financial stress described above, which results in Eq. (3):

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10 The policy-neutral rate is typically defined as the sum of the real equilibrium rate and expected inflation.

11 An explicit definition of an inflation target exists only for countries with an inflation-targeting (IT) regime. Most empirical studies assume, in line with Taylor (1993), that this target does not vary over time and can be omitted from the empirical model.

$$r_t = (1-\rho)[\alpha + \beta(\pi_{t+i} - \pi^*_{t+i}) + \gamma y_t] + \rho r_{t-1} + \delta x_{t+k} + \varepsilon_t \tag{3}$$

Note that the financial stress index x_{t+k} does not appear within the square brackets. This is because it is typically not included in the loss function of central banks' monetary policy but it is rather a factor such as the lagged interest rate, i.e., it may explain why the actual interest rate r_t deviates from the target. Moreover, by placing it in the regression at the same level as a lagged interest rate, we can directly test whether this variable representing ad-hoc policy decisions decreases the interest-rate inertia ϱ, as suggested by Mishkin (2009). At the same time, the response on the coefficient δ can increase, as central banks are more likely to react to financial stress when stress is high. Consequently, it is possible that ϱ and δ move in opposite directions because the central bank either smoothes the interest-rate changes or adjusts the rates in the face of financial stress. In the latter case, the response is likely to be quick and substantial. We set i equal to 6 and k equal to -1.[12] Consequently, the disturbance term ε_t is a combination of forecast errors and is thus orthogonal to all information available at time t (Ω_t).

The empirical studies on monetary-policy rules have moved from using time-invariant estimates (Clarida et al., 1998) through sub-sample analysis (Taylor, 1999; Clarida et al., 2000) toward more complex methods that allow an assessment of the evolution of the conduct of monetary policy. There are two alternative methods for modeling structural changes in monetary-policy rules that occur on an unknown date: (i) regime-switching models, in particular state-dependent Markov switching models (Valente, 2003; Assenmacher-Wesche, 2006; Sims and Zha, 2006) and (ii) state-space models, where the changes are characterized by smooth transitions rather than abrupt switches (Boivin, 2006; Kim and Nelson, 2006; Trecroci and Vassalli, 2009). As argued in Baxa et al. (2010), we consider the second approach to be preferable for the estimation of policy rules, given that it is more flexible and allows for the incorporation of a simple correction of endogeneity (Kim, 2006; Kim and Nelson, 2006), which is a major issue in forward-looking policy rules estimated from ex-post data.[13] The state-space approach, or time-varying coefficient model, also seems suitable when one wants to evaluate the effect of factors such as financial stress that can, for a limited length of time, alter (rather than permanently change) monetary-policy conduct.

State-space models are commonly estimated by means of a maximum likelihood estimator via

12 More precisely, i equals 6 when we use monthly data and 2 for quarterly data. Although the targeting horizon of central banks is usually somewhat longer (4–8 quarters), as in the other papers in this stream of literature, we prefer to proxy inflation expectations by inflation in $t + 2$ quarters for the following reasons. First, the endogeneity correction requires a strong correlation between the endogenous regressor and its instruments. Second, the prediction error logically increases at longer horizons. Most importantly, the choice of i is in line with the theory. Batini and Nelson (2001) show that $i = 2$ in their baseline model of an optimal policy horizon. However, alternative specifications of their model show some sensitivity in terms of what is the optimal i. Nevertheless, employing different i's for regression results left the results in most cases unchanged, to a large extent. In the case of the output gap, we instead assume a backward-looking reaction. The reason is that in the absence of real-time data, we have to rely on the output-gap construction by statistical methods such as HP filter. It is arguable that aside from the prediction error, there is also a construction error that might be magnified if an unobserved forecast is substituted by the output-gap estimate for future periods. Finally, we assume that central bankers' response (if any) to financial stress is rather immediate (see Mishkin, 2009). Therefore, we use one lag of the FSI and its subcomponents in the benchmark case. However, as a robustness check, we allow for different lags and leads, allowing the central bankers' response to financial stress to be preemptive rather than reactive.

13 The time-varying parameter model with specific treatment of endogeneity is still relevant when real-time data are used (Orphanides, 2001). The real-time forecast is not derived under the assumption that nominal interest rates will remain constant within the forecasting horizon (Boivin, 2006) or in the case of measurement error and heteroscedasticity (Kim et al., 2006).

the Kalman filter or smoother. Unfortunately, this approach has several limitations that can become problematic in applied work. First, the results are somewhat sensitive to the initial values of the parameters, which are usually unknown, especially in the case of variables whose impacts on the dependent variable are not permanent and whose sizes are unknown, which is the case for financial stress and its effect on interest rates. Second, the log likelihood function is highly non-linear, and in some cases optimization algorithms fail to minimize the negative of the log likelihood. In particular, it can either fail to calculate the Hessian matrix throughout the iteration process, or, when the likelihood function is approximated to facilitate computations, the covariance matrix of observation vectors can become singular for the starting values provided. The alternative is a moment-based estimator proposed by Schlicht (1981, 2005) and Schlicht and Ludsteck (2006), which is employed in our paper and briefly described below. This framework is sufficiently flexible such that it incorporates the endogeneity correction proposed by Kim (2006).

Kim (2006) shows that the conventional time-varying parameter model delivers inconsistent estimates when explanatory variables are correlated with the disturbance term and proposes an estimator of the time-varying coefficient model with endogenous regressors. Endogeneity may arise not only in forward-looking policy rules based on ex-post data (Kim and Nelson, 2006; Baxa et al., 2010) but also in the case of variables that have a two-sided relationship with monetary policy. Financial stress unquestionably enters this category. Following Kim (2006), we rewrite Eq. 3 as follows:

$$r_t = (1-\rho_t)[\alpha_t+\beta_t(\pi_{t+i})+\gamma_t y_t]+\rho_t r_{t-1}+\delta_t x_{t+k}+\varepsilon_t \quad , \tag{4}$$

$$\alpha_t = \alpha_{t-1}+v_{1,t} \ , \ v_{1,t} \sim i.i.d.\, N(0,\sigma^2_{v_1}) \quad , \tag{5}$$

$$\beta_t = \beta_{t-1}+v_{2,t} \ , \ v_{2,t} \sim i.i.d.\, N(0,\sigma^2_{v_2}) \quad , \tag{6}$$

$$\gamma_t = \gamma_{t-1}+v_{3,t} \ , \ v_{3,t} \sim i.i.d.\, N(0,\sigma^2_{v_3}) \quad , \tag{7}$$

$$\delta_t = \delta_{t-1}+v_{4,t} \ , \ v_{4,t} \sim i.i.d.\, N(0,\sigma^2_{v_4}) \quad , \tag{8}$$

$$\rho_t = \rho_{t-1}+v_{5,t} \ , \ v_{5,t} \sim i.i.d.\, N(0,\sigma^2_{v_5}) \quad , \tag{9}$$

$$\pi_{t+i} = Z'_{t-1}\xi+\sigma_\phi\phi_t \ , \ \phi_t \sim i.i.d.\, N(0,1) \quad , \tag{10}$$

$$y_t = Z'_{t-1}\zeta+\sigma_\psi\psi_t \ , \ \psi_t \sim i.i.d.\, N(0,1) \quad , \tag{11}$$

$$x_{t+k} = Z'_{t-1}\varsigma+\sigma_\iota\iota_t \ , \ \iota_t \sim i.i.d.\, N(0,1) \quad , \tag{12}$$

The measurement Eq. (4) of the state-space representation is the monetary-policy rule. The transitions in Eqs. (5)–(9) describe the time-varying coefficients as a random-walk process without drift.[14] Eqs. (10)–(12) track the relationship between the potentially endogenous regressors (π_{t+i}, y_{t+j}, and x_{t+k}) and their instruments, Z_t. We use the following instruments: π_{t-1}, π_{t-12}, (π_{t-4}, for CAN and SWE), y_{t-1}, y_{t-2}, r_{t-1} and r^f_t, the foreign interest rate for countries other than the United States (the three-month EURIBOR for SWE and UK, and the US three-month interbank rate for CAN and AUS). Unlike Kim (2006), we assume that the parameters in Eqs. (10)–(12) are time-invariant. The correlation between the standardized residuals φ_t, ψ_t, ι_t and the error term ε_t is $\kappa_{\varphi,\varepsilon}$ and $\kappa_{\psi,\varepsilon}$, $\kappa_{\iota,\varepsilon}$, respectively (note that σ_φ, σ_ψ and σ_ι are the standard errors of φ_t, ψ_t, and ι_t, respectively). Consistent

14 Note that while a typical time-invariant regression assumes that $a_t = a_{t-1}$, in this case, it is assumed that $E[a_t] = a_{t-1}$.

estimates of the coefficients in Eq. (4) are obtained in two steps. In the first step, we estimate Eqs. (10)–(12) and save the standardized residuals φ_t, ψ_t, and ι_t. In the second step, we estimate Eq. (13) below along with Eqs. (5)–(9). Note that Eq. (13) now includes bias correction terms, i.e., the (standardized) residuals from Eqs. (10)–(12), to address the aforementioned endogeneity of the regressors. Consequently, the estimated parameters in Eq. (13) are consistent, as u_t is uncorrelated with the regressors.

$$\begin{aligned} r_t &= (1-\rho_t)[\alpha_t+\beta_t(\pi_{t+i})+\gamma_t y_t]+\rho_t r_{t-1}+\delta_t x_{t+k}+\kappa_{\phi,\varepsilon}\sigma_{\phi,\varepsilon}\phi_t+\kappa_{\psi,\varepsilon}\sigma_{\psi,\varepsilon}\psi_t+\kappa_{\iota,\varepsilon}\sigma_{\iota,\varepsilon}\iota_t+u_t\ , \\ u_t &\sim N(0,(1-\kappa^2_{\phi,\varepsilon}+\kappa^2_{\psi,\varepsilon}+\kappa^2_{\iota,\varepsilon})\sigma^2_{\varepsilon,t}) \end{aligned} \tag{13}$$

As previously noted, instead of the standard framework for second-step estimation, the maximum likelihood estimator via the Kalman filter (Kim, 2006), we use an alternative estimation framework, the "varying coefficients" (VC) method (Schlicht, 1981; Schlicht, 2005; Schlicht and Ludsteck, 2006). This method is a generalization of the ordinary least squares approach that, instead of minimizing the sum of the squares of the residual $\sum_{t=1}^{T} u_t^2$ s , uses minimization of the weighted sum of the squares:

$$\sum_{t=1}^{T} u_t^2+\theta_1\sum_{t=1}^{T} v_{1,t}^2+\theta_2\sum_{t=1}^{T} v_{2,t}^2+\ldots+\theta_n\sum_{t=1}^{T} v_{n,t}^2 \tag{14}$$

where the weights θ_i are the inverse variance ratios of the regression residuals u_t and the shocks in time-varying coefficients v_t , that is $\theta_i=\sigma^2/\sigma_i^2$. Hence it balances the fit of the model and the parameter stability.[15.] Additionally, the time averages of the regression coefficients estimated by such weighted least squares estimator are identical to the GLS estimates of the corresponding regression with fixed coefficients, that is, $\frac{1}{T}\sum_{t=1}^{T}\hat{a}_t=\hat{a}_{GLS}$.[16] The method is useful in our case because:

- it does not require knowledge of initial values even for non-stationary variables prior to the estimation procedure. Instead, both the variance ratios and the coefficients are estimated simultaneously;
- the property of the estimator that the time averages of the estimated time-varying coefficients are equal to its time-invariant counterparts, permits easy interpretation of the results in relation to time-invariant results;
- it coincides with the MLE estimator via the Kalman filter if the time series are sufficiently long and if the variance ratios are properly estimated.[17] However, this method suffers from

15 It should be noted that throughout our computations we did not have to solve problems with convergence of the moment estimator, as it was almost always able to find equilibrium. Computational details of the VC method are described in the Appendix. Originally, Schlicht and Ludsteck (2006) start with a derivation of the maximum likelihood estimator of parameters a based on the idea of orthogonal parameterization, which is described in the Appendix. Then they prove that the weighted least squares estimator is identical to the maximum likelihood estimator and also that the likelihood estimator is identical to the moment estimator for very large samples.

16 See Schlicht and Ludsteck (2006) and Baxa et al. (2010) for more details.

17 The Kalman filter as implemented in common econometric packages typically uses the diffusion of priors for its initiation, but it still produces many corner solutions and often does not achieve convergence. Schlicht and Ludsteck (2006) compare the performance of the moment estimator and the Kalman smoother in terms of the mean squared error on simulated data, and they conclude that the moment estimator outperforms the Kalman filter on small

certain limitations of its own. In particular it requires that: (a) the time-varying coefficients are described as random walks, and (b) the shocks in time-varying coefficients v_t are minimized (see Eq. (14)).

While this does not represent a major problem for the estimation of the coefficients of common variables such as inflation, where the monetary-policy response is permanent, it can lead to a loss of some information about ad-hoc response factors in monetary policy making that are considered by central bankers only infrequently; however, once they are in place, the policy response can be substantial. The financial stress indicator x_{t+k} seems to be this kind of factor. One way to address this problem is by estimation-independent calibration of the variance ratios in Eq. (14), such that the estimated coefficient is consistent with economic logic, i.e., it is mostly insignificant and can become significant (with no prior restriction on its sign) during periods of financial stress, i.e., when the financial stress indicator is different from zero. Therefore, we first estimate Eq. (13) using the VC method and study whether the resulting coefficients in the FSI correspond to economic intuition, especially whether the coefficient is not constant or slowly moving (the so-called pile-up problem, see Stock and Watson, 1998). When this problem occurs, we compare the results with models where k belongs to (-2, -1, 0, 1, 2) and calibrate the variance ratios in Eq. (13) by the variance ratios estimated for the model with the largest variances in the FSI. This step was necessary for Australia and Sweden. The Taylor-rule coefficients were compared with the initial estimates and were consistent in both cases.[18]

The results of our empirical analysis should reveal whether central banks adjusted their interest-rate policies in the face of financial stress. However, the time-varying framework also allows for inferring whether any response to financial stress led to the temporal dismissal of other targets, in particular the inflation rate. Therefore, we are mainly interested in the evolution of the financial-stress coefficient δ_t. We expect it to be mostly insignificant or zero, given that episodes of financial stress are rather infrequent, and even if they occur, the monetary authorities may not always respond to them. Moreover, the size of the estimated coefficient does not have any obvious interpretation because the FSI is a composite indicator normalized to have a zero mean. Consequently, we define the stress effect as a product of the estimated coefficient δ_t and the value of the IMF's FSI x_{t+k}. The interpretation of the stress effect is straightforward: it shows the magnitude of interest-rate reactions to financial stress in percentage points or, in other words, the deviation from the target interest rate, as implied by the macroeconomic variables, due to the response to financial stress.

3.4 Results

This section summarizes our results on the effect of financial stress on interest-rate setting. First, the results on the effect of the overall measure of financial stress on interest-rate setting are presented.

samples with a size of up to 100 observations. For comparison, we estimated Eq. (12) using the conventional Kalman filter in the GROCER software using the tvp function (Dubois-Michaux, 2009). We parameterized the model by initial conditions taken from the OLS estimates of the parameters on the full sample and the initial forecast error covariance matrix set to 0. The matrix of the residuals of time-varying coefficients is assumed to be diagonal, as in the VC method. The results were very similar to those obtained from the VC method when the estimated variances were the same in both methods.

18 Stock and Watson (1998) propose a medium-unbiased estimator for variance in the time-varying parameter model, but its application is straightforward only in the case of one time-varying coefficient, and more importantly, it requires the variables to be stationary.

Second, the effect of specific components of financial stress on monetary policy is examined. Third, we briefly comment on the monetary-policy rule estimates that served as the input for the assessment of financial-stress effects. Finally, we perform a series of robustness checks.

3.4.1 Financial-stress effect

Figure 2 presents our results on the effect of financial stress on interest-rate setting in all five countries (referred to as the *financial-stress effect* hereinafter).[19] Although there is some heterogeneity across countries, some global trends in the effect of financial stress are apparent. Whereas in good times, such as in the second half of the 1990s, financial stress has virtually no effect on interest-rate setting or is slightly positive,[20] the reaction of monetary authorities to financial stress was highly negative during the 2008–2009 global financial crisis. While the previous evidence on the effect of financial stress on monetary policy is somewhat limited, our results broadly confirm the time-invariant findings of Cecchetti and Li (2008), who show that the US Fed adjusted interest rates to the procyclical impact of bank capital requirements in 1989–2000. Similarly, Belke and Klose (2010) estimate the Taylor rule on two sub-samples (before and during the 2008–2009 global financial crisis) and find that the Fed reacted systematically not only to inflation and the output gap, but also to asset prices, credit, and money.

The size of financial-stress effects on interest-rate setting during the recent financial crisis is somewhat heterogeneous, with the strongest reaction found for the UK. The results suggest that all central banks except the Bank of England maintain policy rates at approximately 50–100 basis points lower compared to the counterfactual policy of no reaction to financial stress. The size of this effect for the UK is assessed to be approximately three times stronger (i.e., 250 basis points). This implies that approximately 50% of the overall policy-rate decrease during the recent financial crisis was motivated by financial-stability concerns in the UK (10%–30% in the remaining sample countries), while the remaining half falls to unfavorable developments in domestic economic activity. This finding complements previous results suggesting that the BoE's consideration of expected inflation over the last decade has been very low (as found by Baxa et al., 2010, using the time-varying model and by Taylor and Davradakis, 2006, in the context of the threshold model) by evidence that it further decreased during the current crisis. It is also evident that the magnitude of the response is unusual for all five central banks. However, the results for Australia, Canada, and Sweden show a similar magnitude of response to financial stress during the recent financial crisis compared to that observed in previous periods of high financial stress.

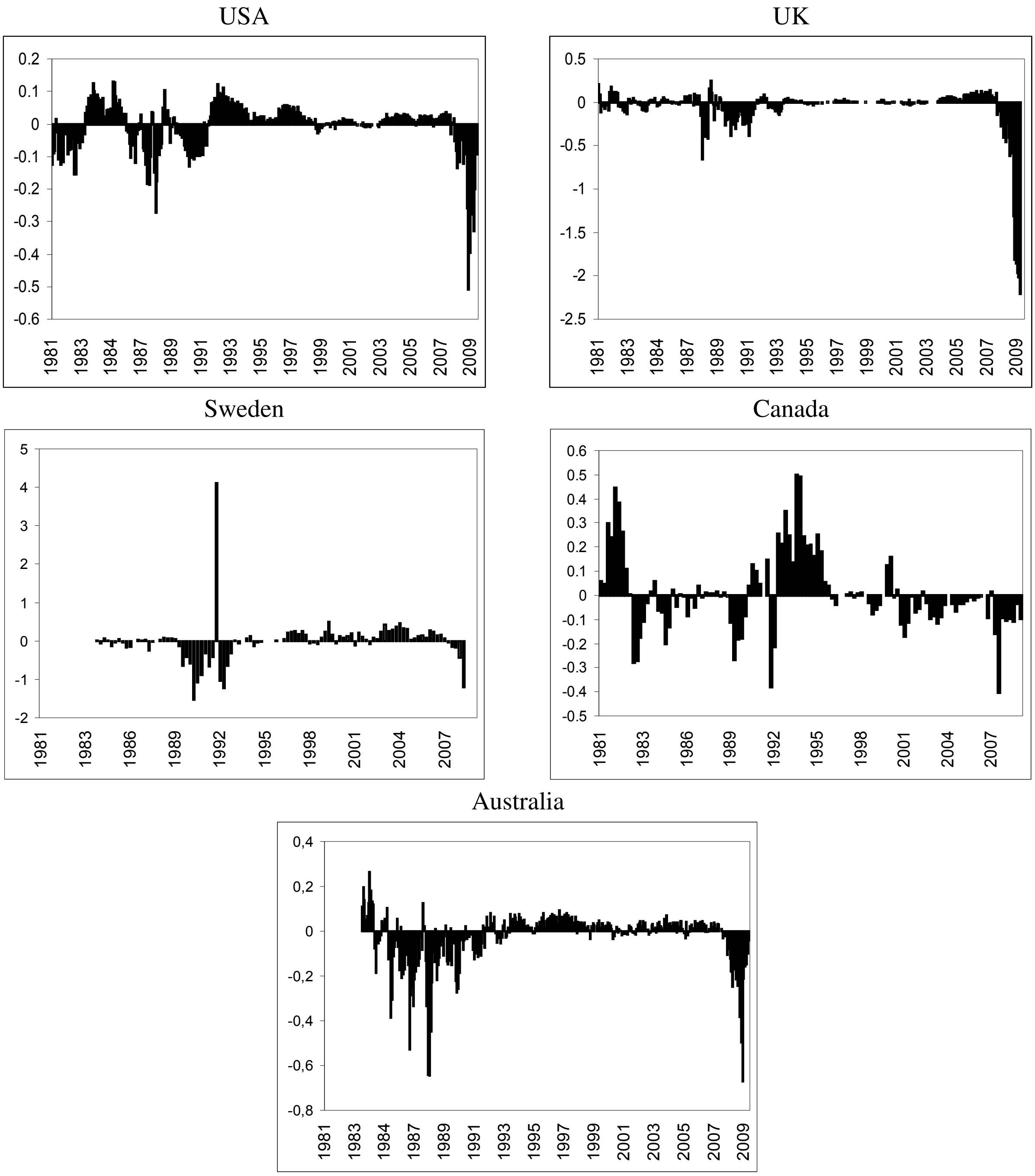
Given that the 2008–2009 global crisis occurred at the end of our sample (there is a peak in the stress indicator of five standard deviations that has not returned to normal values yet), we performed an additional check to avoid possible end-point bias. In particular, we ran our estimation excluding the observation from the period of the 2008–2009 crisis. These results were practically indistinguishable from the full sample estimation. With regard to the effect of the current crisis, the largest uncertainty is associated with the results for Canada, for which the shortest data sample – ending in the fourth quarter of 2008 – was available. When the possibility of a pre-emptive reaction

19 Given that the magnitude of the financial-stress effect differs across countries, especially due to the high positive peak for Sweden and negative peak for the UK, we use different scales for different countries.

20 Note that the positive effect of financial stress on interest-rate setting is to some extent a consequence of scaling the financial-stress indicator; its zero value corresponds to the long-run average stress. Hence, we do not pay much attention to positive values of stress unless caused by a temporarily positive and significant regression coefficient associated with the FSI.

of the central bank to financial stress is considered (see the robustness checks below), the effect of financial stress in the current crisis is estimated for Canada at somewhere between 1% and 2% (see Appendix 3). These additional results suggest that the response of the Bank of Canada in the benchmark model is likely to be underestimated.

Figure 2 – The Effect of Financial Stress on Interest-Rate Setting



Notes: The figure depicts the evolution of the financial-stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the financial-stress indicator in the monetary-policy rule and the value of the IMF financial-stress indicator (δx). The stress effect shows the magnitude of the interest-rate reaction to financial stress in percentage points.

The question of which components of financial stress influence interest-rate setting is addressed in Figure 3. In this case, we estimate the model using each FSI subcomponent separately (the bank stress effect, the exchange-rate stress effect, and the stock-market stress effect) instead of the overall FSI and report the financial-stress effect attributable to each subcomponent. Some heterogeneity across countries is again apparent, although it seems that bank stress and stock-market stress dominated central bankers' considerations in less open economies. On the other hand, exchange-rate stress matters in more open economies such as Canada and Sweden.

Specifically, the US Fed seemed to be worried about financial instability, especially during the 1980s. We can observe that the main concern in the early 1980s was banking stress, which is arguably related to the Savings and Loans crisis. Another concern was that of stock-market stress, in particular during the stock-market crash of 1987, when interest rates were 30 b.p. lower with respect to the benchmark case.

The Bank of England was, in general, much more perceptive to financial stress. We find its response mainly to stock-market stress again, notably, in 1987. Interestingly, we find little response to exchange-rate stress, not even during the 1992 ERM crisis. Nevertheless, it has to be emphasized that the interest-rate reaction to this speculative attack was subdued in comparison to, for example, the Riksbank (Buiter et al., 1998). The base rate was increased by 2 p.p. to 12% on September 16, 1992. Despite a promise of further increases up to 15%, traders continued selling the pound. On the evening of the same day, the UK left the ERM with interest rates unchanged; on the following day, the base rate decreased to 10.5%; and at the end of September, the base rate was 9%, lower than at the beginning of the month. Therefore, despite huge open market operations, the response of the interest rate was moderate, with the monthly interest-rate average practically unaffected. Hence, our framework does not detect any effect of financial stress on the interest rate during the ERM crisis. Since the devaluation of the pound sterling in September 1992, the effect of financial stress on interest-rate setting approaches zero from originally negative values. Aside from this, the response of the Bank of England to inflation has decreased. From this perspective, it seems the pound sterling's withdrawal from the ERM allowed for both a more rule-based and less restrictive monetary policy. With respect to the banking crisis in the late 2000s, the Bank of England provided liquidity support in its earlier stage in 2007 with the fall of Northern Rock. Policy rates remained constant until late 2008, despite the bankruptcy of Lehman Brothers in the US in September 2008. The reason for keeping policy rates constant was related to concerns regarding potential inflationary pressures from rising oil and food prices.

The interest-rate effect of the banking crisis in Sweden in the early 1990s is estimated to be slightly over 1% in absolute terms (see Figure 2). The crisis began in September 1990, when the non-banking financial institution Nyckeln unexpectedly collapsed (Jennergren, 2002). The Riksbank did not decrease interest rates sharply because coincidental international factors, in particular the reunification of Germany, forced interest rates upwards. Despite facing recession, the government attempted to defend the peg of the krona to ECU and decided to prevent the spread of the banking crisis by announcing a blanket guarantee for the liabilities of the banking sector (Jonung, 2009). Hence, interest-rate cuts were not a primary tool chosen for resolution of the crisis.

In comparison to the United Kingdom, the reaction of the Riksbank to the ERM crisis was different. First, after a series of speculative attacks on the Swedish krona in mid-September 1992, the Riksbank still attempted to maintain the fixed exchange rate, and the marginal interest rate jumped up 500% to offset the outflow of liquidity and other speculative attacks (see the large positive stress effect on the interest rate in 1992 in Figure 2). However, not even such an increase was sufficient, and the fixed exchange rate had to be abandoned later, in November.[21]

The Reserve Bank of Australia significantly loosened its policy during the 1980s. This can be attributed to stress in the banking sector with the exception of the reaction to the stock-market crash in 1987 (see Figure 3).

The exchange rate as well as bank stress seems to matter for interest-rate considerations at the Bank of Canada. Interestingly, the results suggest that the Bank of Canada often responded to higher exchange-rate stress by monetary tightening. A possible explanation for this finding might be that given the openness of the Canadian economy, its central bank tightened the policy when the currency stabilized at the level that the monetary authority considered to be undervalued.

We would like to highlight a comparison of Figures 2 and 3. First, it should be noted that a positive response to one stress subcomponent may cancel out in the face of a negative response to another one, making the response to the overall stress negligible (as in the case of Canada). Second, the stress effects related to individual subcomponents do not necessarily sum up to the stress effect related to the entire FSI.

Overall, the results suggest that the central bank tends to react to financial stress, and different components of financial stress matter in different time periods. The effect of financial stress on interest-rate setting is found to be virtually zero in good times and economically sizable during periods of high financial stress.

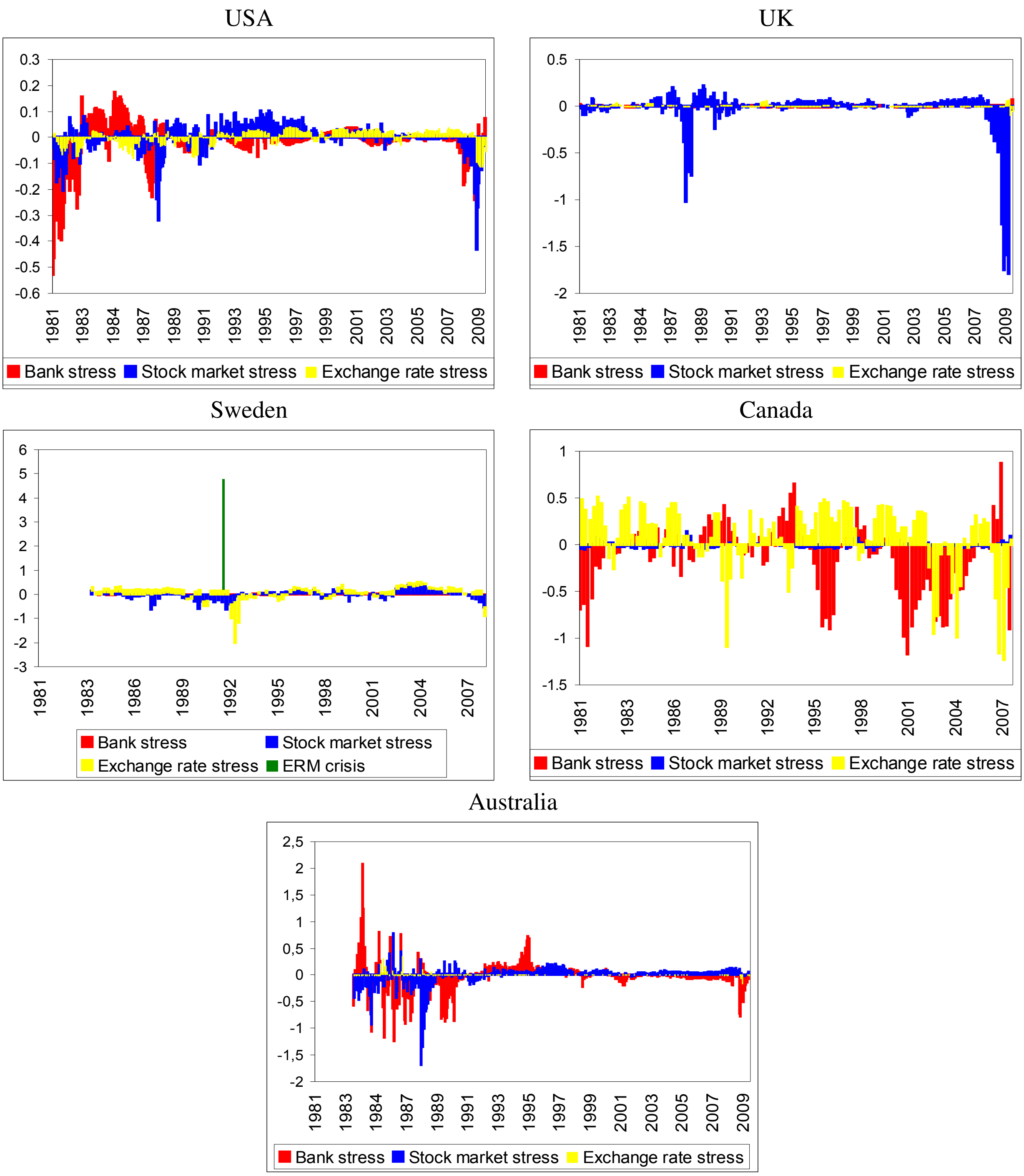
3.4.2 Monetary policy rule estimates

Given that our main interest lies in the interest-rate response to financial stress, we comment on the other monetary-policy rule estimates only briefly. The plot of the evolution of the estimated parameters over time for all countries is available in Appendix 1. First of all, it should be noted that most coefficients do indeed vary over time, which is consistent with previous evidence and underlines the fact that monetary-policy conduct has evolved substantially in recent decades.

In general, the responses to inflation (β) are positive, and the coefficient is often above one, consistent with the Taylor principle. Nevertheless, we find that in the last decade the coefficient decreased somewhat, and during the recent financial crisis it even turned slightly negative (in the US and UK; more on this below). The decrease of the inflation response during the last decade is typically attributed to well-anchored inflation expectations as well as a low-inflation environment (Sekine and Teranishi, 2008; Baxa et al., 2010). The finding of negative β during the recent crisis is likely to be related to the fact that central banks were decreasing policy rates to historical lows in the face of exceptionally high financial stress, despite inflation expectations being largely unchanged, rather than being an indication that policy rates were systematically decreased when inflation expectations increased.

21 For Sweden, we add a dummy variable for the third quarter of 1992 (ERM crisis) to Eq. 13. At this time, the Swedish central bank forced short-term interest rates upward in an effort to keep the krona within the ERM. From the perspective of our model, it was a case of a strong positive reaction to the actual stress that lasted only one period. When this dummy variable was not included, the model with a lagged value of the FSI was unable to show any link between stress and interest rates, and the estimates of other coefficients were inconsistent with economic intuition. Clearly, since we use data at monthly and quarterly frequency, this limits the possibility to detect and properly analyze day-to-day dynamics of some short-term instability events.

Figure 3 – The Effect of Financial Stress Components on Interest-Rate Setting: Bank Stress, Exchange-Rate Stress, and Stock-Market Stress



Notes: The figure depicts the evolution of the components of the financial-stress effect, namely, the bank-stress effect, the exchange-rate stress effect, and the stock-market stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the given component of the financial-stress indicator in the monetary-policy rule and the value of the corresponding component of the IMF financial-stress indicator (δx). The stress effect shows the magnitude of the interest-rate reaction to financial stress in percentage points.

For the United States, our results show that the response to inflation was highest in the early 1980s, and except for the period following the recession of 1990–1991 the estimated coefficient is higher or very close to one. This value is slightly lower in comparison to Kim and Nelson (2006), who found the response to be around 1.5 and almost invariant since 1981. Given the size of the confidence intervals, it is, however, difficult to determine whether our results differ significantly. Kim and Nelson (2006) estimate the interest-rate smoothing coefficient to be higher than 0.8, i.e., in line with what time-invariant estimates of monetary-policy rules typically suggest (see, for example, Clarida et al., 1998, 2000). Our estimates indicate that the interest-rate smoothing is somewhat lower (0.5–0.6). This finding is in line with the recent critique by Rudebusch (2006), who argues that the practical unpredictability of interest-rate changes over a few quarters suggests that the degree of interest-rate smoothing is rather low. Interestingly, we find that the response to inflation decreases substantially after the terrorist attacks on September 11, 2001. This complies with Greenspan (2007), who argued in that case that the Fed was concerned about the US economy spiraling downward into recession after the terrorist attacks. Later, Greenspan himself acknowledged that the monetary policy was somewhat loose, but *ex ante* optimal, given the increased uncertainty after the attacks. In a similar vein, Taylor (2010) compares the actual values of the federal funds rate and the counterfactual values predicted by the (time-invariant) Taylor rule, finding that in 2002–2005 interest rates were too low compared to predictions and this deviation from a rules-based policy was "larger than in any period since the unstable decade before the Great Moderation" (p. 167). Negative estimates of the response to inflation in this particular period are reported also by Trecroci and Vassalli (2010). The response to the output gap is significant for nearly the whole sample, although the values close to 0.2 are somewhat lower than in Kim-Nelson (2007), but similar to Trecroci and Vassalli (2010).

The results for countries that currently have an explicit target for inflation share several features. The interest-rate smoothing is again found to be lower in comparison to time-invariant estimates, with midpoints around 0.5. The exception is Canada, where the values fluctuate around zero and are insignificant. Moreover, for some central banks, such as the RBA and the BoE in 2010 or the Sveriges Riksbank in late 1980, we find that central banks are less inertial during crises.[22] Second, the response of interest rates to inflation is particularly strong during the periods when central bankers want to break a record of high inflation, such as in the UK or Australia at the beginning of the 1980s, and is less aggressive in a low-inflation environment with subdued shocks and well-anchored inflation expectations (Kuttner and Posen, 1999). In this respect, our results confirm the findings of Taylor and Davradakis (2006), who argue that the response of the Bank of England to inflation is insignificant when the inflation rate is close to its target. Third, some central banks (Australia and Canada) are also found to react to output-gap developments, with the parameter estimated to be slightly positive on average, whereas the parameter is insignificant with wide confidence intervals in Sweden and the United Kingdom.

The results show that the interest-rate response to financial stress is insignificant most of the time, at the 95% significance level. This is in line with our expectations, i.e., that the coefficients should be insignificant in periods when stress is low. Nevertheless, the coefficient on financial stress is statistically significant at the 95% level during the recent financial crises for most countries. The importance of financial stress for interest-rate setting is further confirmed using the

22 Indeed, the correlation coefficient of the estimated time-varying coefficient of the lagged interest rate ρ and the financial-stress index δ is -0.79 for Australia, 0.21 for Canada, -0.20 for Sweden, -0.68 for the UK, and 0.60 for the US.

GMM estimation, which shows that the financial-stress index is significant, in fact, in all countries. In addition, when the one-standard-deviation quantile is taken into account instead of the more usual two-standard-deviation quantile, the periods when we can identify any interest-rate response to financial stress become more evident. We present a list of these periods in Table A6.2 in the Appendix.

3.4.3 Robustness checks

In terms of the financial-stress effect estimates, we perform a battery of robustness checks. First, following the argument put forward above that the interbank rate may occasionally provide a better signal of monetary-policy intentions than the policy rate, we use interbank interest rates as a dependent variable. These results are reported in Figures A.2.1–A.2.2. We can observe that the overall stress effect on the interbank rate was larger for the US during the current crisis, where it explains 2% of the decrease of the interbank interest rate. For Sweden, we found a strong positive effect of exchange rate volatility in the late 1980s; this might be linked to the aim of the central bank to keep the exchange rate fixed. In other cases, there is no substantial difference between the benchmark results and the results obtained using this alternative dependent variable.

Second, in the benchmark model and all of the results reported thus far, we use the first lag of the FSI in the policy-rule estimation. We motivate this choice by the use of monthly data, the frequency of monetary-policy meetings of most central-bank boards, and the assumption that policy actions are likely to be implemented in a timely fashion. In addition, we employ different lags and leads, in the latter case allowing the policy to be preemptive rather than reactive. In this case, we use the future realized value of the FSI as a proxy for the central bank's expectation (in a similar manner as to how it is routinely executed for inflation expectations) and, consequently, treat the FSI as an endogenous variable (see Figure A.3.1 for the results). To obtain comparable results, we calibrate the variance ratios with the same values as in the baseline specification. Although we find rather mixed evidence on preemptive policy actions, which may also be related to the inadequacy of proxying the expected values of financial stress by the actual values of the financial-stress indicator as well as the fact that a central bank might not react to the stress preemptively, the reaction to financial stress in the current crisis is strongly negative for both expected and observed stress.

Third, we further break down the FSI sub-indices to each underlying variable to evaluate their individual contributions.[23] The corresponding stress effects appear in Figures A.4.1–A.4.2. Breaking down stock-market-related stress, we find that the US Fed and the BoC react to the corporate bond spread, whereas the BoE and Sveriges Riksbank are more concerned with stock returns and volatility. While the RBA seems to be concerned with both corporate bond spreads and stock-market volatility in the 1980s, the role of stock-related stress had substantially decreased by then. As far as bank-related stress is concerned, the TED spread plays a major role in all countries apart from the UK, where the largest proportion of the effect on the interest rate can be attributed to an inverted term structure.

Fourth, because the verifications related to comparing our econometric framework to obvious alternatives such as, first, the use of a maximum likelihood estimator via the Kalman filter instead of the moment-based time-varying coefficient framework of Schlicht and, second, the use of a Markov switching model instead of a state-space model, were provided in Baxa et al. (2010), we

23 This applies only to the banking and stock-market subcomponents because the foreign-exchange subcomponent is represented by a single variable.

estimate simple time-invariant monetary-policy rules for each country by the generalized method of moments, including various subsamples. This simple evidence reaffirms that the analyzed central banks seem to pay attention to overall financial stress in the economy. The FSI is statistically significant, with a negative sign and a magnitude of between 0.05–0.20 for all countries. On the other hand, the coefficients of its subcomponents often are not significant, and the exchange-rate subcomponent in some cases has a positive sign. These results, which are available upon request, confirm that to understand the interest-rate adjustment in response to financial stress, one should rely on a model allowing for a differential response across time.

3.5 Conclusion

The 2008–2009 global financial crisis generated significant interest in exploring the interactions between monetary policy and financial stability. This paper aimed to examine in a systematic manner whether and how the monetary policy of selected main central banks (the US Fed, the Bank of England, the Reserve Bank of Australia, the Bank of Canada, and Sveriges Riksbank) responded to episodes of financial stress over the last three decades. Instead of using individual alternative measures of financial stress in different markets, we employed the comprehensive indicator of financial stress recently developed by the International Monetary Fund, which tracks overall financial stress as well as its main subcomponents, in particular banking stress, stock-market stress and exchange-rate stress.

Unlike a few existing empirical contributions that aim to evaluate the impact of financial-stability concerns on monetary policy making, we adopt a more flexible methodology that not only allows for the response to financial stress (and other macroeconomic variables) to change over time, but also addresses potential endogeneity (Kim and Nelson, 2006). The main advantage of this framework is that it not only enables testing of whether central banks responded to financial stress at all, but also detects the periods and types of stress that were the most worrying for monetary authorities. Our results indicate that central banks truly change their policy stances in the face of financial stress, but the magnitude of such responses varies substantially over time. As expected, the impact of financial stress on interest-rate setting is essentially zero most of the time, when the levels of stress are very moderate. However, most central banks loosen monetary policy when the economy faces high financial stress. There is some cross-country and time heterogeneity when we examine central banks' considerations of specific types of financial stress. While most central banks seem to respond to stock-market stress and bank stress, exchange-rate stress is found to drive the reaction of central banks only in more open economies

Consistent with our expectations, the results indicate that a sizeable fraction of the monetary-policy easing during the 2008–2009 financial crisis can be explained by a direct response to the financial stress above what might be attributed to the decline in inflation expectations and output below its potential. However, the size of the financial-stress effect differs by country. The result suggests that all central banks except the Bank of England kept their policy rates at 50–100 basis points lower, on average, solely due to the financial stress present in the economy. Interestingly, the size of this effect for the UK is assessed at about three times stronger (i.e., 250 basis points). This implies that about 50% of the overall policy-rate decrease during the recent financial crisis was motivated by financial-stability concerns in the UK (10%–30% in the remaining sample countries), while the remaining half falls to unfavourable developments in domestic economic activity. For the US Fed, macroeconomic developments themselves (a low-inflation environment and output substantially below its potential) explain the majority of the interest-rate policy decreases during the crisis, leaving any further response to financial stress to be constrained by zero interest rates.

Overall, our results point to the usefulness of augmenting the standard version of monetary-policy rules by some measure of financial conditions to obtain a better understanding of the interest-rate-setting process, especially when financial markets are unstable. The empirical results suggest that the central banks considered in this study altered the course of their monetary policy in the face of financial stress. The recent crisis seems truly to be an exceptional period, in the sense that the response to financial instability was substantial and coincided in all the countries analyzed, which is evidently related to intentional policy coordination absent in previous decades. However, we have also observed that previous idiosyncratic episodes of financial distress were, at least in some countries, followed by monetary-policy responses of similar, if not higher, magnitude.

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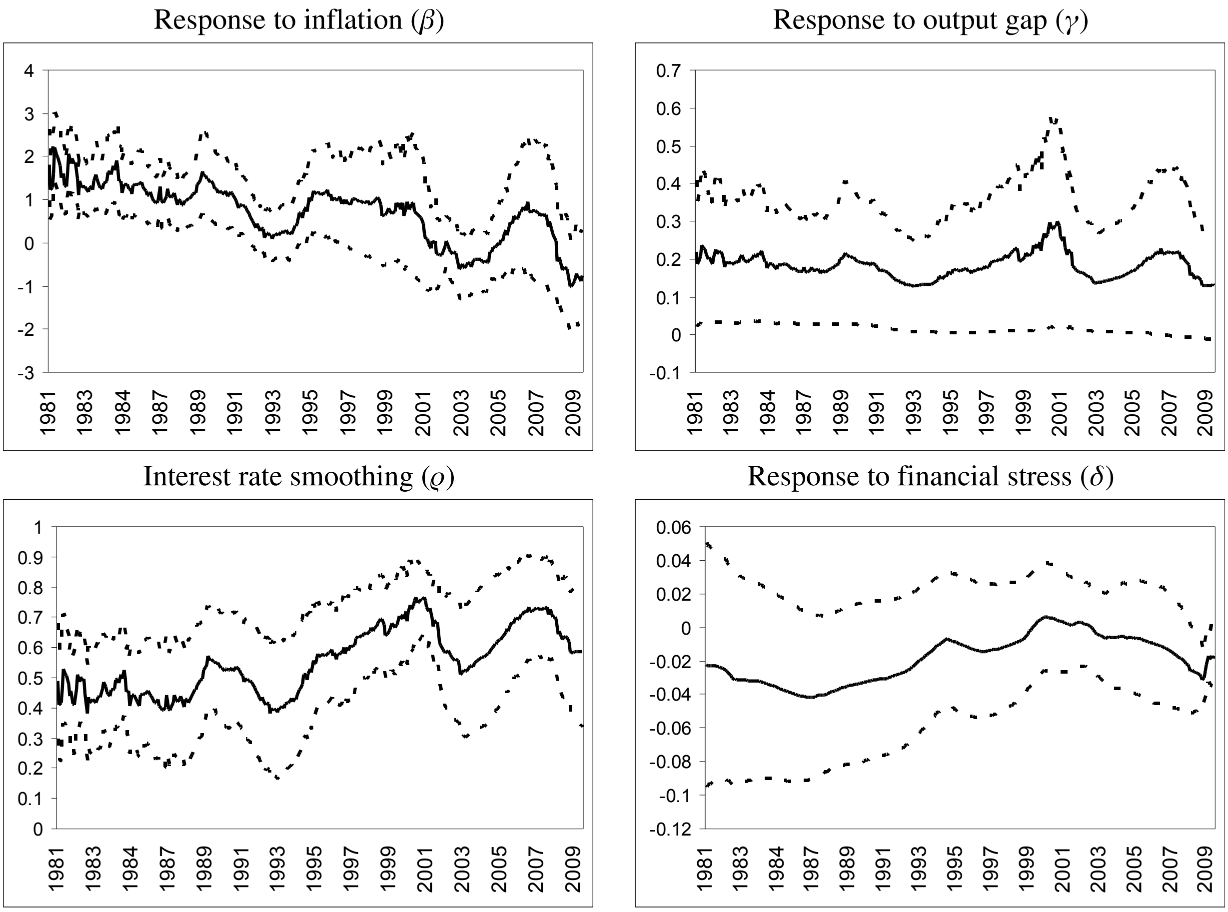
Appendix

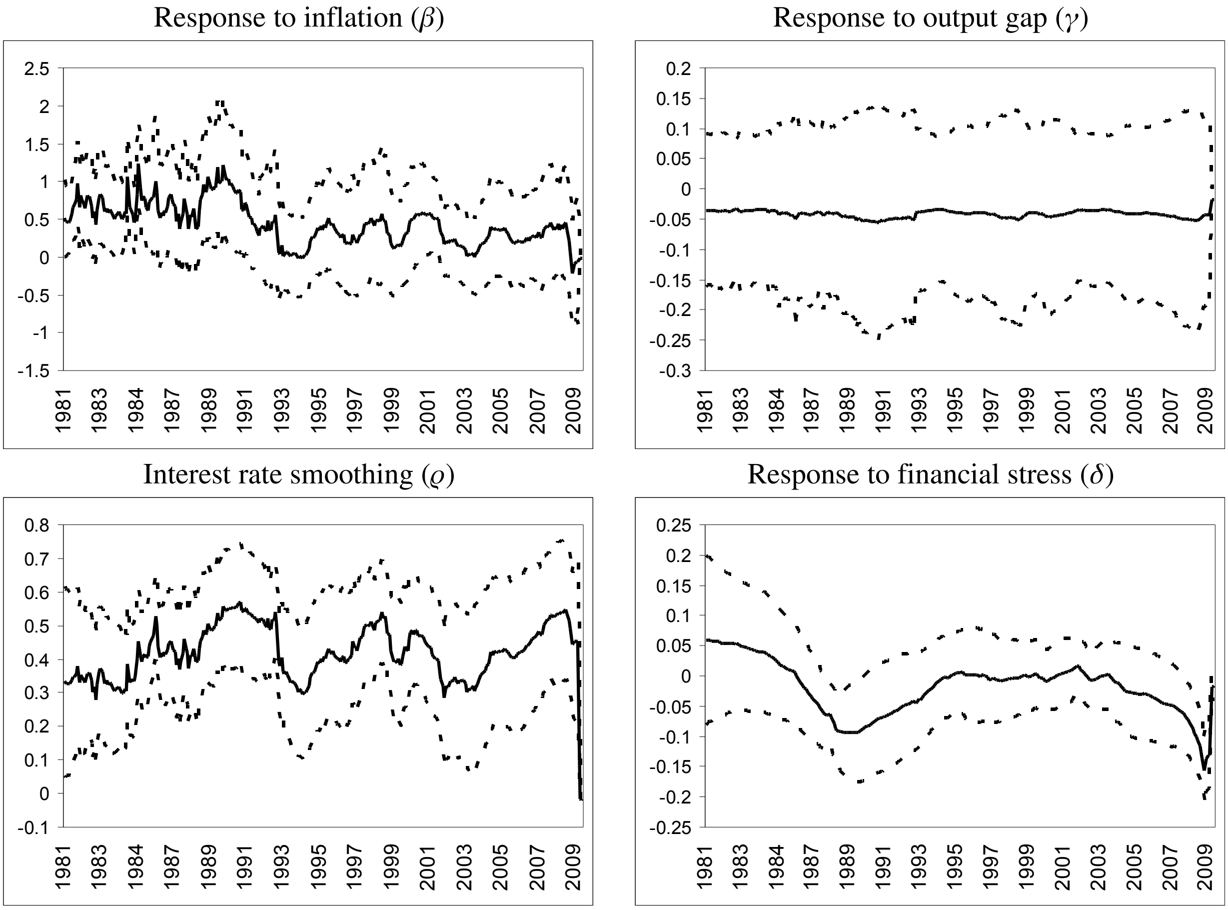
Appendix 1 Time-Varying Monetary Policy Rule Estimates

Figure A.1.1 – Time-Varying Monetary Policy Rules: USA



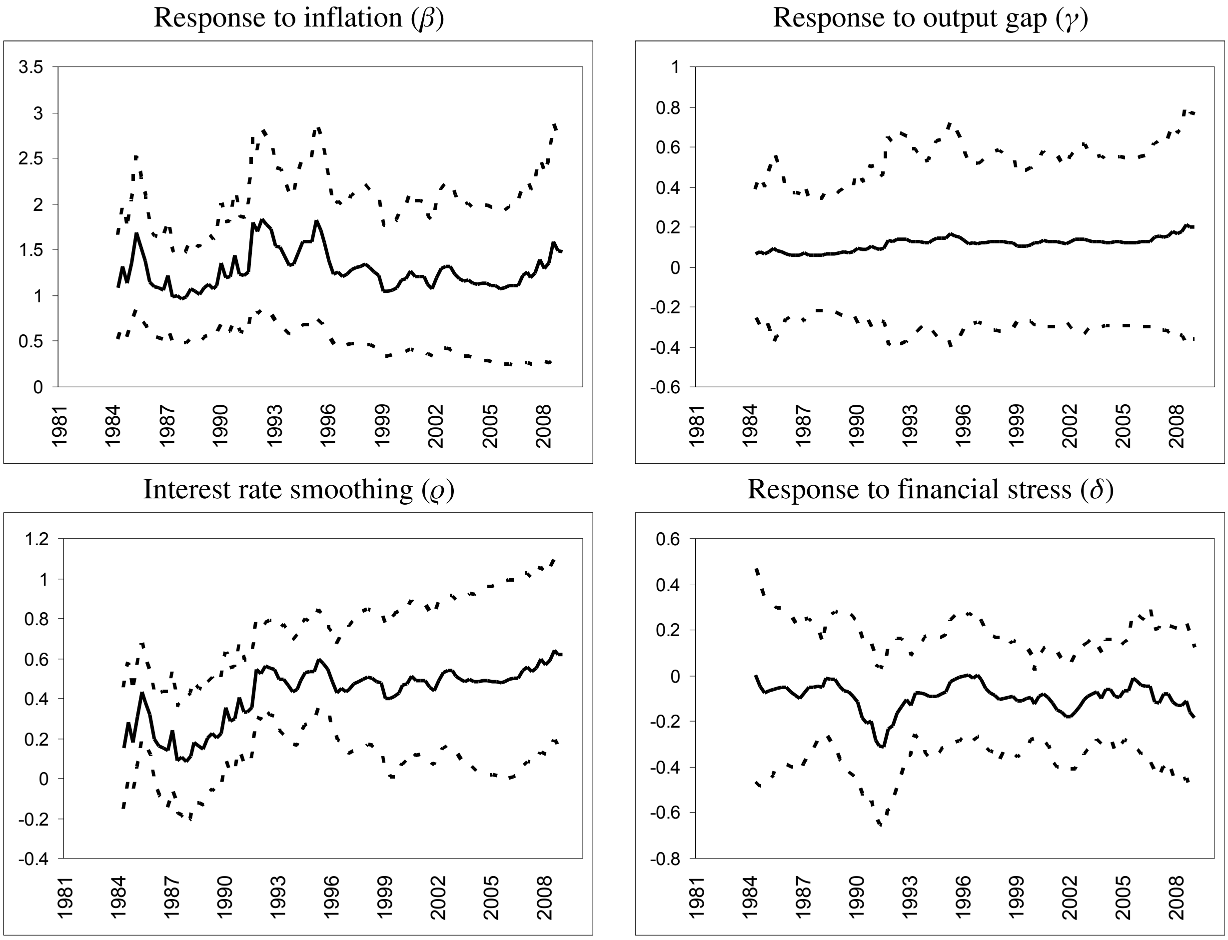
Note: The estimated coefficients of the time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.1.2 – Time-Varying Monetary Policy Rules: UK



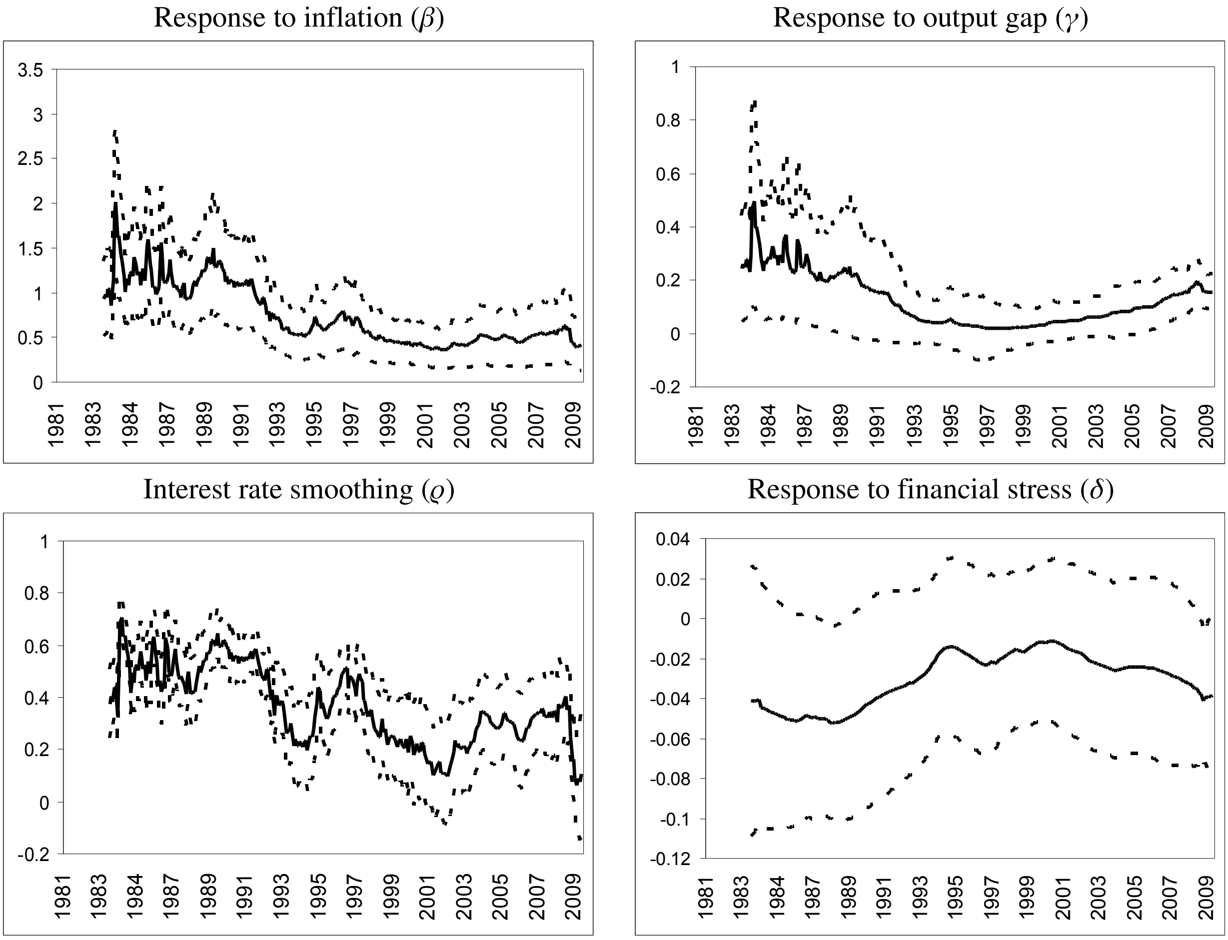
Note: The estimated coefficients of the time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.1.3 – Time-Varying Monetary Policy Rules: Sweden



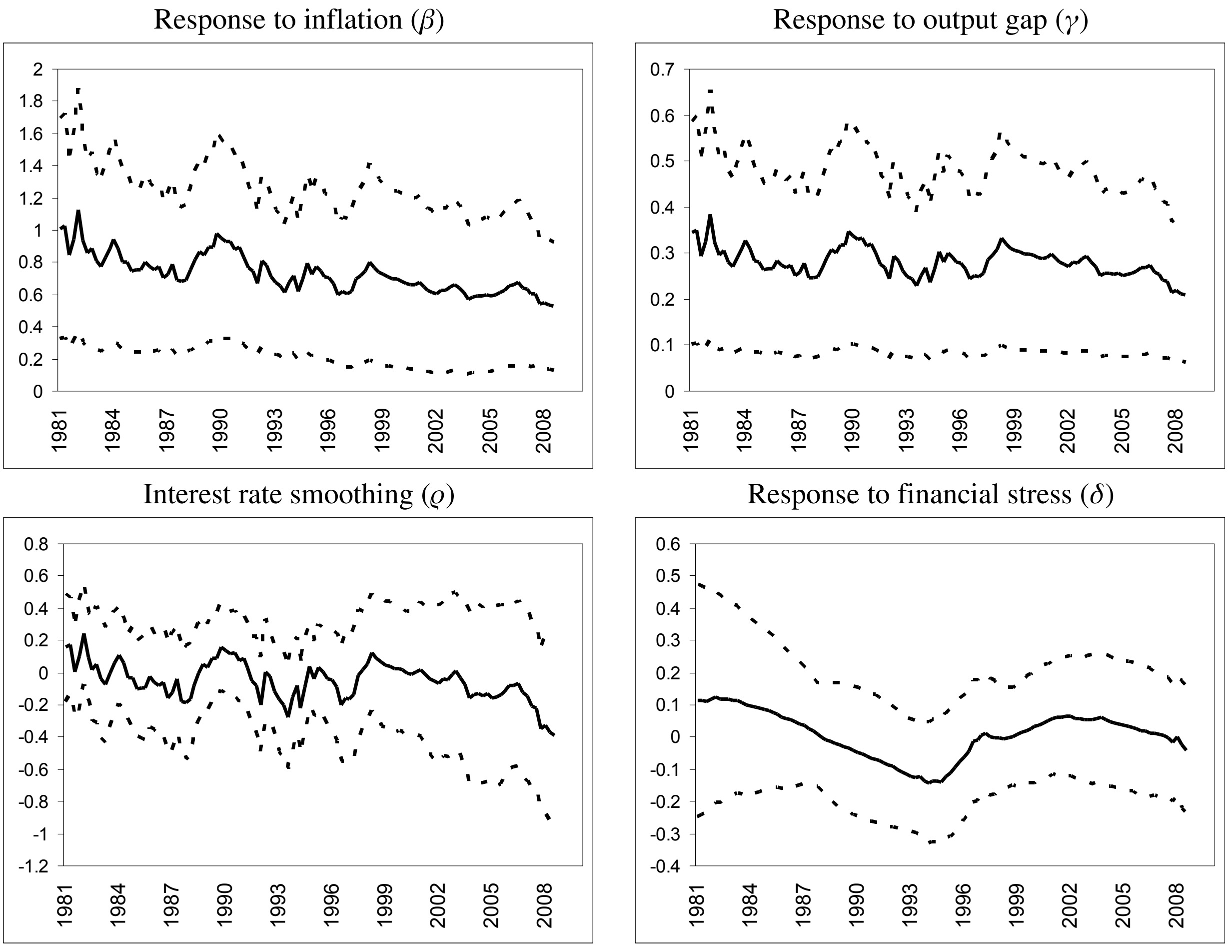
Note: The estimated coefficients of the time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.1.4 – Time-Varying Monetary Policy Rules: Australia



Note: The estimated coefficients of the time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.1.5 – Time-Varying Monetary Policy Rules: Canada

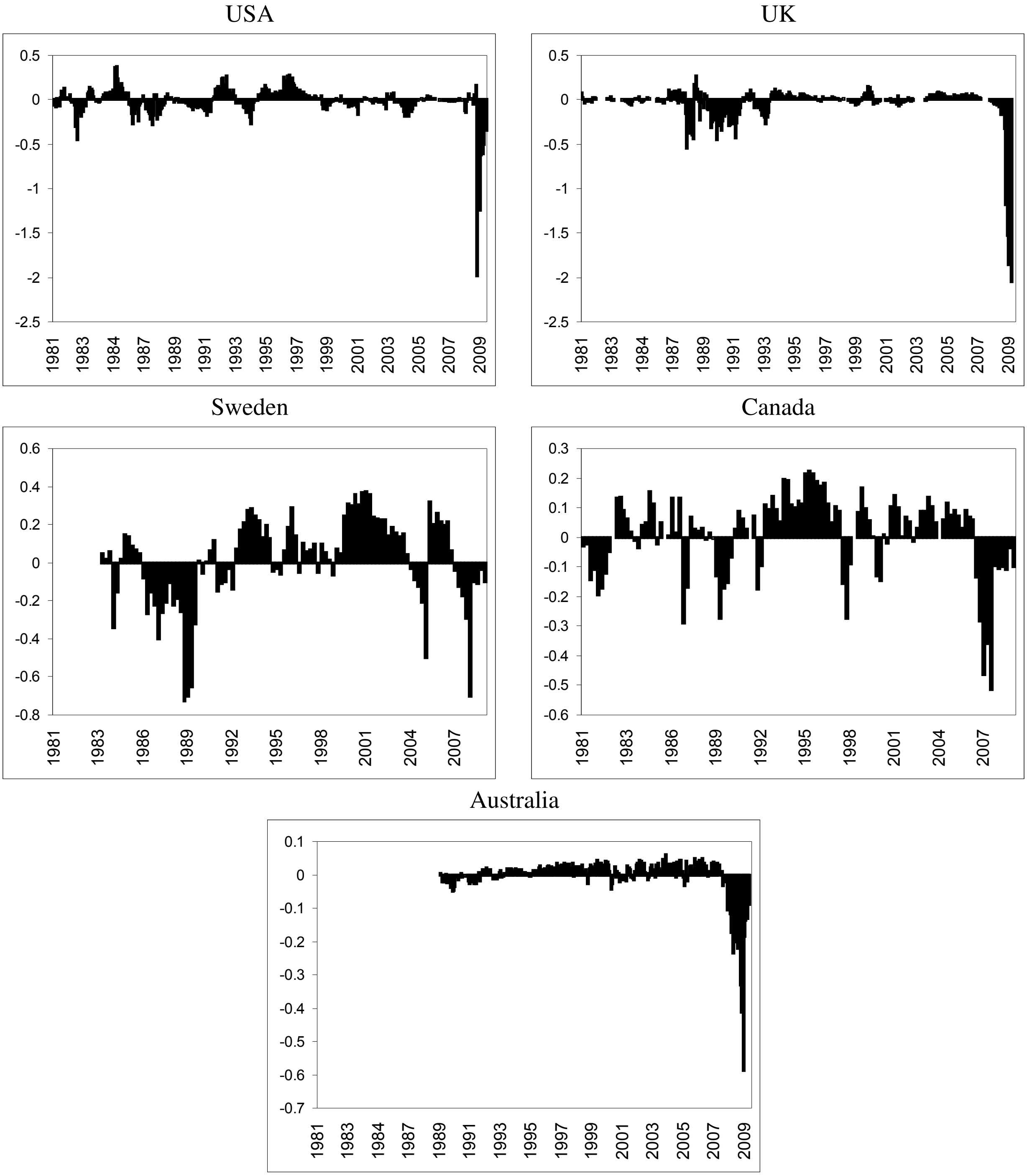


Note: The estimated coefficients of the time-varying monetary policy rule are depicted with a 95% confidence interval.

Appendix 2 The Results with the Interbank Rate in the Policy Rule

Figure A2.1 – The Effect of Financial Stress on Interest-Rate Setting

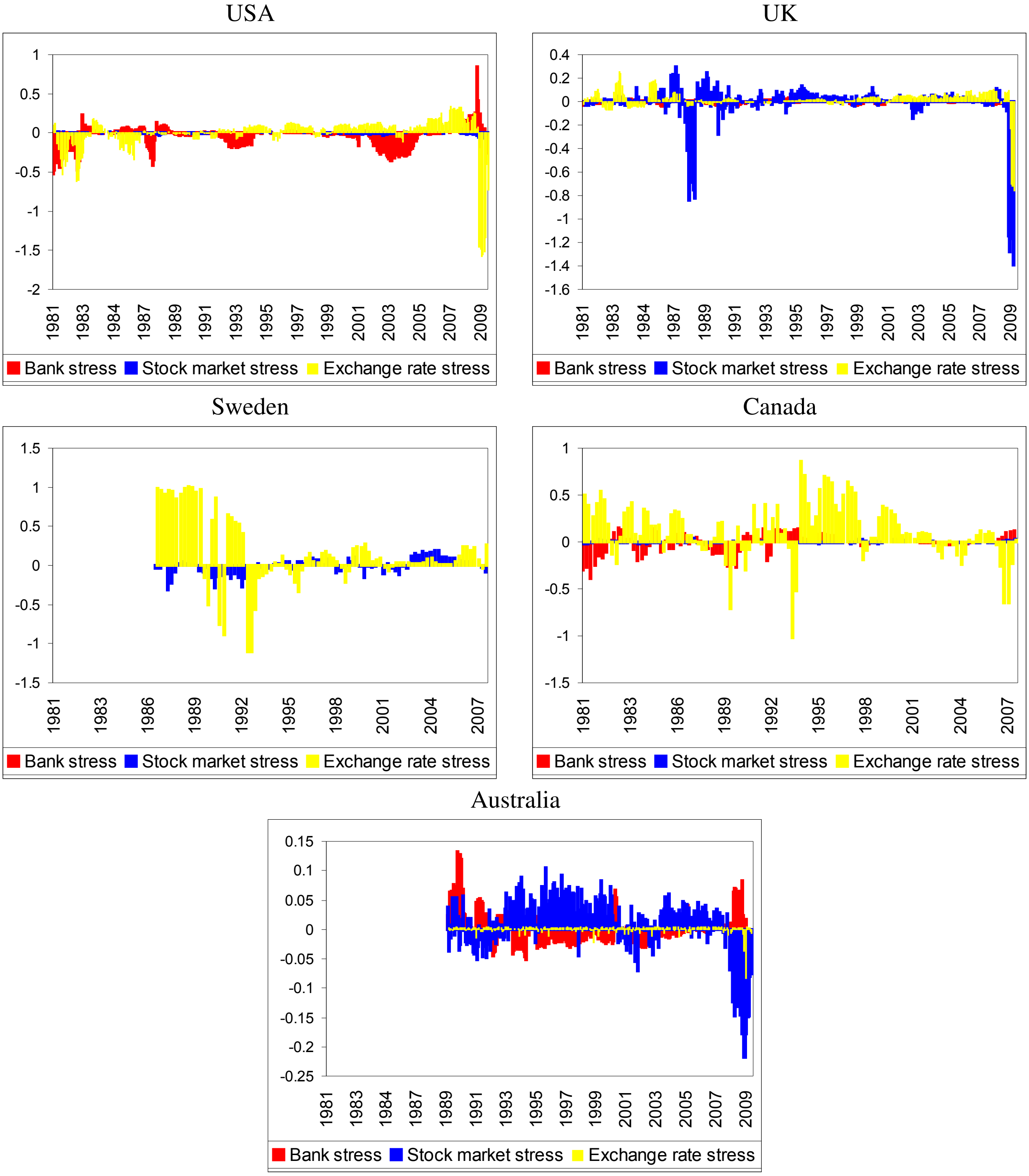
(Interbank interest rate as the dependent variable in the policy rule)



Notes: The figure depicts the evolution of the financial-stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the financial-stress indicator in the monetary-policy rule and the value of the IMF financial-stress indicator (δx). The stress effect shows the magnitude of the interest-rate reaction to financial stress in percentage points.

Figure A2.2 – The Effect of Financial Stress Components on Interest-Rate Setting:

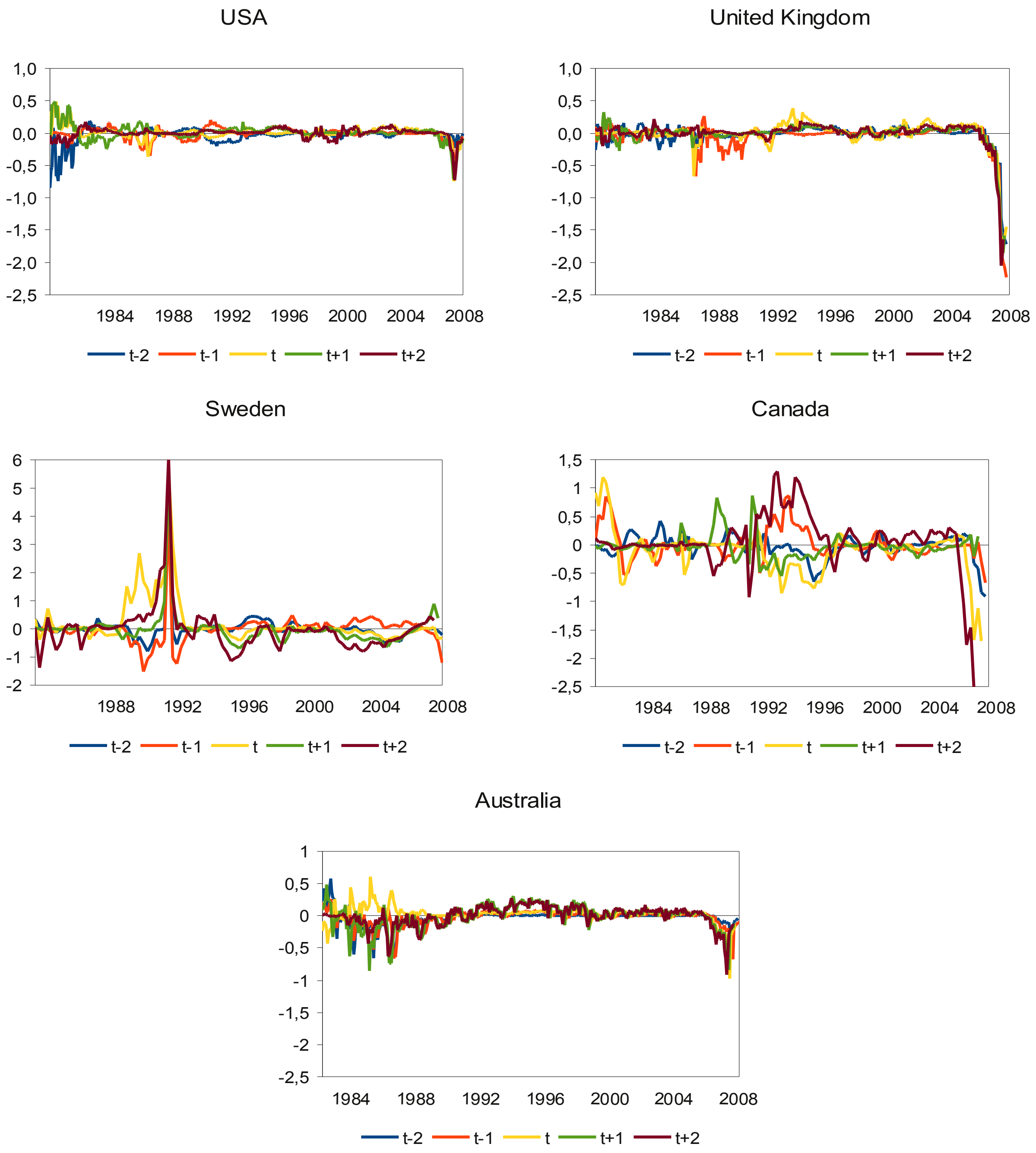
Bank Stress, Exchange-Rate Stress and Stock-Market Stress



Notes: The figure depicts the evolution of the components of the financial-stress effect, namely, the bank stress effect, the exchange-rate stress effect, and the stock-market stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the given component of the financial-stress indicator in the monetary-policy rule and the value of the corresponding component of the IMF financial-stress indicator (δx). The stress effect shows the magnitude of the interest-rate reaction to financial stress in percentage points.

Appendix 3 The Results with Different Leads and Lag of the FSI

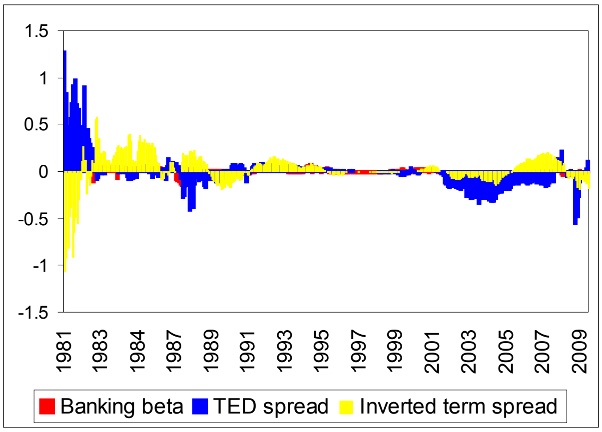
Figure A3.1 – The Effect of Financial Stress (t-1 vs. t-2, t, t+1, t+2) on Interest-Rate Setting



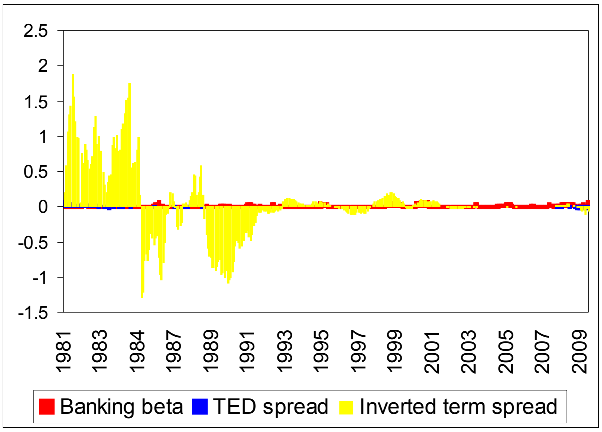
Appendix 4 The Results with Individual Variables of Bank Stress and Stock-Market Stress

Figure A4.1 – The Effect of Bank Stress on Interest-Rate Setting

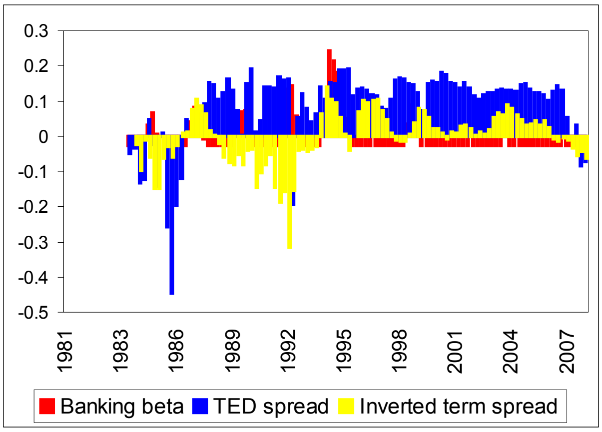
USA



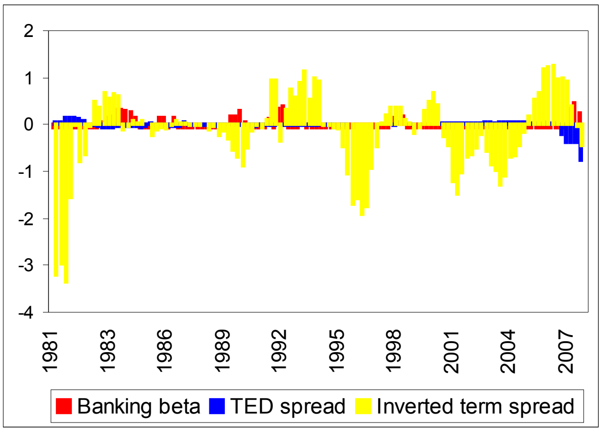
UK



Sweden



Canada



Australia

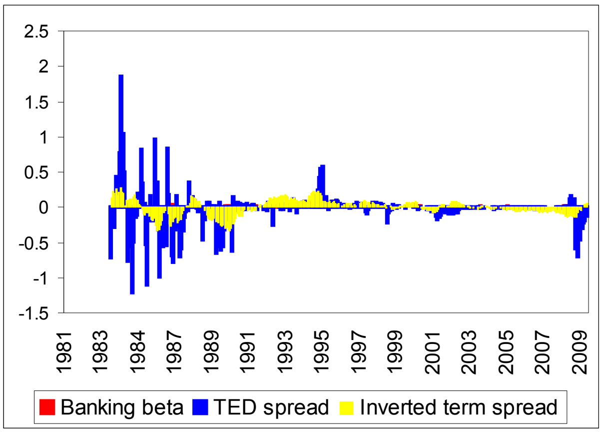
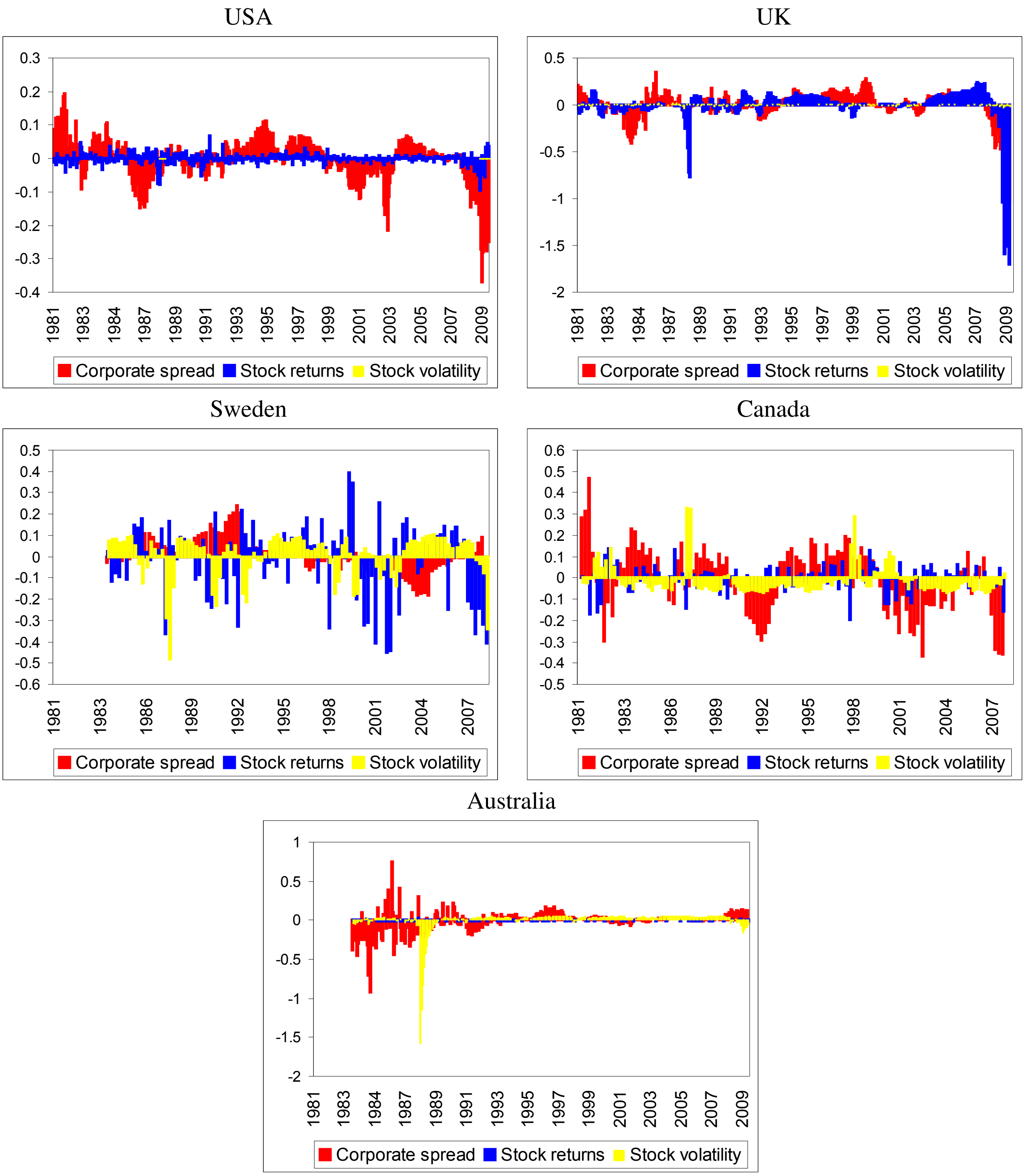


Figure A4.2 – The Effect of Stock-Market Stress on Interest-Rate Setting



Appendix 5 Data

Appendix 6 Tables

Table A6.1: Significance of the Financial Stress Index in the Estimated Taylor Rules
Time-Invariant Reaction Functions, GMM estimates

	α	**β**	**γ**	**ϱ**	**δ**	**J-statistics**	**p-value**
				United States			
1	5.59	1.59	1.51	0.97	-0.014	24.9808	*0.6289*
	(1.42)	*(0.99)*	*(0.53)*	*(0.01)*	*(0.006)*		
2	-9.42	0.45	0.94	0.94		15.0451	*0.8207*
	(4.78)	(0.33)	(0.34)	(0.01)			
				United Kingdom			
1	3.93	0.37	1.51	0.97	-0.018	15.0423	*0.5212*
	(1.31)	*(0.52)*	*(0.39)*	*(0.01)*	*(0.004)*		
2	7.68	-0.89	2.77	0.98		11.4534	*0.4905*
	(2.78)	*(0.74)*	*(0.71)*	*(0.01)*			
				Sweden			
1	-1.87	2.59	-0.16	0.84	-0.135	24.9808	*0.6289*
	(0.86)	*(0.46)*	*(0.22)*	*(0.04)*	*(0.029)*		
2	0.24	2.03	-0.12	0.76		15.0451	*0.8207*
	(0.53)	*(0.39)*	*(0.16)*	*(0.05)*			
				Australia			
1	0.04	2.06	-0.02	0.95	-0.038	21.5261	*0.9731*
	(0.79)	*(0.3)*	*(0.1)*	*(0.01)*	*(0.006)*		
2	2.2	1.84	0.19	0.89		15.2464	*0.9830*
	(0.93)	*(0.22)*	*(0.14)*	*(0.02)*			
				Canada			
1	-0.33	2.07	0.87	0.89	-0.089	10.6859	*0.8284*
	(1.23)	*(0.96)*	*(0.30)*	*(0.04)*	*(0.023)*		
2	1.21	1.67	0.75	0.86		9.4332	*0.7395*
	(1.23)	*(0.87)*	*(0.28)*	*(0.05)*			
				United States: 1981:1–1999:12 sample			
1	1.82	1.43	1.48	0.95	-0.015	20.1672	*0.8583*
	(1.27)	*(0.57)*	*(0.27)*	*(0.01)*	*(0.007)*		
1*	-0.25	2.18	0.3	0.87	-0.043	19.8946	*0.8683*
	(0.77)	*(0.42)*	*(0.06)*	*(0.02)*	*(0.012)*		
				United States: Clarida et al. (1998) – 1982:10–1994:12 sample			
2	-0.1	1.83	0.56	0.97		10.9000	*0.9980*
	(1.54)	*(0.45)*	*(0.16)*	*(0.03)*			

Notes: Numbers in (.) are standard errors. The samples are as follows: United States: 1981:1M–2009:6M, United Kingdom: 1981:1M–2009:3M, Australia: 1983:3M–2009:5M, Sweden 1984:2Q–20091Q, Canada 1981:1Q–2008:4Q. Model 1: $r_t = (1\text{-}\varrho)(\alpha + \beta\pi_{t+k} + \gamma y_t) + \varrho r_{t-1} + \delta x_{t-1}$. Model 2 does not contain financial stress: $r_t = (1\text{-}\varrho)(\alpha + \beta\pi_{t+k} + \gamma y_t) + \varrho r_{t-1}$. k equals 6 for the USA, the UK and Australia, 2 for Sweden and 4 for Canada. For Sweden, a dummy variable for the third quarter of 1992 (the RM crisis) is included.

Both models are estimated using the GMM. The list of instruments follows. United States: lags of interest rate, output gap, inflation, and financial stress (1, 2, 3, 4, 5, 6, 9, 12), model 2 without lags of FSI in a set of instruments. United Kingdom: lags of interest rate, output gap, inflation, and EURIBOR 3M (1, 2, 3, 4, 5, 6, 9, 12), FSI (1,2,3). Australia: interest rate, inflation, output gap, US money market rate and FSI (1, 2, 3, 4, 5, 6, 9, 12). Sweden: lags of interest rate, inflation, output gap, EURIBOR 3M, and FSI (1, 2, 3, 4) + the dummy for the ERM crisis. Canada: interest rate, inflation, output gap, U.S. money market rate, and FSI (1, 2, 3, 4).

Additionally, we show the results for the USA estimated on the subsample 1981–1999. The model denoted as 1* has the output gap derived from the quadratic trend of log industrial production in a similar fashion as in Clarida et al. (1998). Their results are provided for comparison with ours.

Table A6.2: Periods with Significant Responses to Financial Stress

		1980s	1990s	2000s
United States	2SD			2008:M03-2009:M03
	1SD	1982:M11-1992:M09		2007:M05-2009:M06
United Kingdom	2SD	1987:M08-1989:M11		2007:M09-2009:M03
	1SD	1987:M01-1993:M01		2006:M03-2009:M03
Sweden	2SD			
	1SD		1990:Q2-1992:Q2	2001:Q2-2002:Q3
			1993:Q1	2009:Q1
			1999:Q4-2000:Q2	
Australia	2SD	1987M:04-1988:M10		2008:M09-2009:M03
	1SD	1983:M07-1993:M10		2002:M10-2009:M05
			1996:M06-1997:M05	
Canada	2SD			
	1SD	1982:Q3-1984:Q1	1992:Q3-1995:Q4	

Chapter 4

Fiscal developments and financial stress: A threshold VAR analysis

4.1 Introduction

During periods of economic downturn or stress in financial markets the effects of fiscal developments on economic activity might be different from what is usually observed in good or normal times. Evidence shows that economic downturns are often associated with periods of financial stress or even with financial crisis. Under those circumstances, the share of non-performing loans increases and negative sentiments in the markets depress the value of other financial assets. In some cases, the disruptions in financial markets or problems in the bank balance sheets may trigger a recession by reducing the credit flow to the other sectors. Therefore, it is important to assess the effects of fiscal developments during the periods of market stress to check, notably whether there are non-linearities at play and if fiscal multipliers are different.

Certainly, the relation between financial instability and economic policy can be two-sided. On the one hand, irrespectively of the causes of financial instability, policy makers may try to soften its effect on the economy. On the other hand, so-called "bad" policies can also contribute to financial instability. For instance, a situation of large government indebtedness might cause a loss of confidence in the ability of the government to pay back orderly the outstanding stock of government debt. As a result, unsustainable fiscal policies undermine sovereign debt credibility and financial markets may refuse to buy new government debt and secondary market liquidity may decline. The reduced liquidity can weaken the balance sheet of the banks and other financial institutions that hold government debt. Balance sheet losses related to the price drops in government securities can affect negatively the lending capacities of the banks, which might reduce the credit flow to the private sector. Hence, it is relevant to examine whether and how the effects of fiscal developments on economic activity differ in times of financial instability.

Our contribution to the literature comprises the estimation of the effects of fiscal shocks using a threshold VAR approach, including a measure representing financial instability, namely a Financial Stress Index (Cardarelli et al., 2010). More specifically we use quarterly data, for the U.S., the U.K., Germany and Italy, for the period 1980:4-2009:4, encompassing macro, fiscal and financial variables. Therefore, the use of quarterly fiscal data is another relevant contribution in this context. Moreover, according to our knowledge, there have been no attempts to investigate empirically the effects of fiscal developments associated with periods of financial crises within a multi-equation framework, which is the issue addressed in this paper.

Our main results are as follows: (i) the use of a nonlinear framework with regime switches, determined by a financial stress indicator, is corroborated by nonlinearity tests; (ii) the responses of economic growth to a fiscal shock are mostly positive in both low and high financial stress regimes; (iii) differences in the estimated multipliers across financial regimes are relatively small for the full sample; (iv) the estimated nonlinear impulse responses suggest that the size of the fiscal multipliers is higher than average in the 2008-2009 crisis, except in the U.K.; (v) fiscal policy can attenuate financial stress; (vi) financial stress has a negative effect on output growth and worsens the fiscal situation; (vii) however, the responses to large financial stress shocks are more than proportional especially when the economy is initially in the non-stress regime

The paper is organised as follows. Section two reviews the literature. Section three explains the methodology. Section four describes the data and a review on the fiscal development is provided. Section five conducts the empirical analysis, and section six concludes.

4.2 Related literature

4.2.1 Fiscal VARs

VAR models, in addition to the New Keynesian DSGE models, have become the most popular tool for investigating the effects of monetary policy during the nineties, and a number of stylized facts have been broadly identified. In response to a contractionary shock in the short-term interest rate, (i) real GDP declines with a hump-shape pattern, with a maximum decline occurring between one and one and half year, (ii) the price level declines persistently, and (iii) there is an evidence for a strong liquidity effect, that is, the non-borrowed reserves drop in response to an increase of interest rates. A summary of the research in this field can be found in Christiano, et al. (1999).

However, no such broad consensus has emerged from the research on the effects of fiscal policy, notably regarding the qualitative responses of macroeconomic aggregates to changes in government expenditures or revenues. In this context, the main difficulties come from the approaches used to identify the changes in fiscal policy, since both government expenditures and revenues, to some extent, automatically respond to fluctuations in economic activity and thus these fluctuations need to be distinguished from deliberate policy changes. It is possible to separate these effects using estimated elasticities of tax revenues and government expenditures on output developments or to use external information such as the expected contemporary effects of the fiscal variables. Nevertheless, the differences in the identification schemes in the VAR analysis often lead to different results. For instance, van Brusselen (2010) provides a broad overview of the effectiveness of fiscal policy, and an evaluation of fiscal multipliers notably in several VAR models.

Caldara and Kamps (2008) compared the four existing approaches to identify fiscal policy shocks in VAR models using a dataset for the United States: (i) the Structural Vector Autoregression (SVAR) following Blanchard and Perotti (2002) and Perotti (2005) with calibrated sizes of the automatic stabilizers, (ii) the recursive identification scheme with the Choleski decomposition,[1] (iii) the sign-restriction approach proposed for the analysis of monetary policy by Uhlig (2005) and applied by Mountford and Uhlig (2009), and (iv) the so called "narrative approach" assigning dummy variables associated with periods that are known for exogenous changes in fiscal policy, related to the increases in military build-ups. The authors argue that different identification and calibration schemes lead to similar results as far as the effect of government expenditures is concerned, e.g. the shock to government expenditures is likely to increase output. However, results are rather diverging regarding the responses to changes in taxes.

1 The ordering used in these studies is as follows: government expenditures, G, revenues, T, gross domestic product, Y (all in real per capita terms and natural logs; sometimes the share of G and T on Y is used and they are often augmented for transfers and interest payments), inflation, π (measured as the GDP deflator), and short-term interest rate i.

Romer and Romer (2007) applied a narrative approach in a similar fashion as they did in their 1989 paper on monetary policy. They went through the Congressional records and presidential speeches to identify both timing and size of the changes in taxation. Based on this identification, they find that tax increases were highly contractionary with multipliers that reached the value of three. This value is much higher than the values obtained from other VARs which are concentrated around one. Such discrepancy was explained by Favero and Giavazzi (2009) who argued that the results of Romer and Romer are caused by their estimation method based on one equation. After using the shocks by Romer and Romer within a multivariate framework, Favero and Giavazzi obtained results similar to those from traditional fiscal VARs.

The fiscal VAR approach based either on the SVAR or on the recursive identification was applied for several countries namely in the European Union. Aarle et al. (2003) estimated the effects of fiscal and monetary policy for the members of the European Monetary Union and found significant differences in reactions among the individual countries of the EMU. Muscatelli et al. (2002) found a significant decrease in the responsiveness of the fiscal policy variables in the U.S. since 1979, and similar decreases were also reported for Italy, Germany, France and the United Kingdom.

For Germany, Heppke-Falk et al. (2006), using a VAR approach, mention that government expenditure shocks increase output and private consumption on impact with low statistical significance, while they decrease insignificantly private investment. They also found for government investment – in contrast to government consumption – a positive output effect, which is statistically significant until 12 quarters ahead. In addition, anticipated expenditure shocks have significant effects on output when the shock is realized, but not in the period of anticipation. The authors claim that the effects of expenditure shocks are only short-lived in Germany and government net revenue shocks do not affect output with statistical significance. However, they provide evidence that direct taxes lower output significantly, while small indirect tax revenue shocks have little effect. Moreover, the compensation of public sector employees is equally not effective in stimulating the economy.

For Italy, Giordano et al. (2007), also within a VAR framework, found that a shock to government purchases of goods and services has a sizeable and robust effect on economic activity: an exogenous 1% (in terms of private GDP) shock increases private real GDP by 0.6% after 3 quarters. The response declines to zero after two years, reflecting with a lag the low persistence of the shock. The authors also mention that the effects on employment, private consumption and investment are positive for Italy. In contrast, changes of public sector wages have no significant effect on output, while the effects on employment turn negative after two quarters. Shocks to net revenue were found to have negligible effects on all the variables.

The baseline specification was extended for an analysis of the impact of exchange rate (Monacelli and Perotti, 2006) and for government debt (Favero and Giavazzi, 2007 and Afonso and Sousa, 2009). Afonso and Sousa (2009a, b) used quarterly fiscal data from the U.S., the U.K., Germany and Italy along with the feedback from government debt, and also included the effects on asset markets in a Bayesian VAR model.

For instance, Afonso and Sousa (2009b) using a Bayesian SVAR model provide some evidence that the government spending shocks have, *inter alia*, in general a small effect on GDP; do not impact significantly on private consumption and have a negative effect on private investment in the

U.S., U.K., Germany and Italy. On the contrary; they found that government revenue shocks: have a positive (although) lagged effects on GDP and private investment. Interestingly, they found that when the debt dynamics is explicitly taken into account, (long-term) interest rates and GDP become more responsive and the effects of fiscal policy on these variables also become more persistent. Moreover, the results from Afonso and Sousa (2009b) also provide weak evidence of stabilizing effects of the debt level on the primary budget balance. They also find that government spending shocks, in general, have a positive, but small effect on GDP and also uncover a crowding out effect, which is present in all four countries.

Kirchner, Cimadomo and Hauptmeier (2010) use time-varying structural VAR techniques in the euro area for the period 1980-2008. They report that the short-run effectiveness of government spending in stabilizing real GDP and private consumption has increased until the end-1980s but it has decreased thereafter, and that government spending multipliers at longer horizons have declined substantially.

Regarding the possibility of negative fiscal spending multipliers, and the so-called non-Keynesian effects of fiscal policy, several authors have argued along those lines. For instance, Alesina and Perotti (1996), Giavazzi and Pagano (1998, 2005), and Mitra (2006) mention that high government debt implies additional fiscal stress and a higher probability of higher taxes in the future. Therefore, higher private savings may arise and lower output, and thus the effects of increased government expenditure on output might be negative. In addition, there is also some evidence of expansionary fiscal contractions, the most prominent examples are Denmark in 1993-1985 and Ireland in 1985-1988, and Rzonca and Cizkowitz (2005) identified a similar pattern in the Central and Eastern European countries that have entered the EU in 2004-2006. However, Afonso (2010) reports that the empirical evidence for the EU countries is quite diverse in this respect, notably with alternative definitions of fiscal consolidation episodes.

4.2.2 Fiscal policy and financial instability

Fiscal policy can contribute to financial instability if, for instance, the issuance of substantial amounts of sovereign debt causes fiscal stress and a potential fiscal and/or financial crisis. In particular, unsustainable fiscal policies may undermine sovereign debt credibility and financial markets may refuse to buy new government debt, while transactions in the secondary market may also become less frequent. The inability to sell government bonds reduces its liquidity and weakens the balance sheet of the banks and of other financial institutions that hold government debt. The balance sheet losses related to the price drops in government debt securities negatively affect the lending capacities of the banks, which consequently might reduce the flow of credit to the private sector. Moreover, some related discussion drawing on the fiscal theory of price level (Leeper, 1991, Sims, 1994, and Woodford, 1994, 1995), and its application to Krugman's model of financial crisis (1979) as introduced in Daniel (2001) and Corsetti and Mackowiak (2006) also highlights such possible links.

The effects of fiscal policy can differ in times of financial instability. This links with the Keynesian-like story about countercyclical economic policy, and the possible positive impacts of fiscal stimuli. The idea is that the government steps in to compensate the decline in private sector demand in order to stabilize aggregate demand. Almunia et al. (2009), who compared the policies during the Great Depression and the 2008-09 crisis concluded that when fiscal policy was used in the 1930s it worked, while the evidence for the effectiveness of the monetary policy is rather mixed.

4.2.3 Fiscal policy and financial instability: empirics

The literature dealing with the effects of fiscal policy during the periods of financial stress is relatively scarce, but growing. Baldacci et al. (2008) tried to answer the question of whether fiscal policy might shorten the recession caused by banking crisis. Using OLS estimation and truncated Logit on a dataset containing 118 banking crises in 99 countries 1980-2000, they have found that fiscal policy responses are significant for the duration of the crisis, and that the composition of the fiscal package is a key to success. In this respect their results are in line with Blanchard et al. (2009) who tried to summarize the policy recommendations from the empirical literature in order to give guidelines for the construction of fiscal stimuli packages that had been prepared at that time.

On the other hand Bouthevillain and Dufrénot (2010) who used a Markov switching model with time-varying probabilities within a single-equation framework have not found such differences in the efficiency of fiscal policy in France. Similarly Afonso, Grüner and Kolerus (2010, using a panel of OECD and non-OECD countries, for the period, could not reject the hypothesis that the effects of fiscal policy are essentially the same in the absence and during a financial crisis

On the other hand, in terms of the effects of monetary policy, there are several papers addressing this issue, namely Balke (2000), Atanasova (2003), Li and St-Amant (2008) and Berkelmans (2005). Berkelmans (2005) included a variable representing credit frictions in a small SVAR model of the Australian economy and has shown that monetary policy might in this case play a stabilizing role and it can reduce the effects of credit shocks on output.

Using a threshold vector autoregression with credit conditions as a threshold variable, Balke (2000) has shown that the U.S. output responds more to monetary policy in a credit-rationed regime. Atanasova (2003) analyzed the impact of credit frictions on business cycles dynamics in the U.K. and her results in many respects confirm the conclusions by Balke (2000). Finally, Li and St-Amant (2008) estimated a threshold vector autoregression for the monetary transmission mechanism in Canada with an indicator of financial stress (Illing-Liu, 2006) as a threshold variable, and have estimated explicitly the nonlinear properties of the system. Their findings indicated that there are nonlinear effects of contractionary and expansionary shocks and that the large contractionary shocks increase the likelihood of moving to high stress regime. Furthermore, the high stress regime is in their dataset typically associated with weaker output growth, higher inflation and higher interest rates. However, and as far as we can tell, there are no studies that investigate empirically the effects of fiscal developments during recessions associated with periods of financial crises within a multi-equation framework, and that is precisely what we do in this paper.

4.3 Methodology

4.3.1 Threshold Vector Autoregression

We follow the approach used by Balke (2000) and Atanasova (2003) for the estimation of a TVAR. Thus, we include a threshold variable in the fiscal VAR, for which we have chosen the financial stress index (FSI), introduced by the IMF (see Cardarelli et al., 2010) and modified by Balakrishnan et al. (2009).

The TVAR model has a number of interesting features that make it attractive for our purposes. First, it is a relatively simple way to capture possible nonlinearities such as asymmetric reactions to shocks or the existence of multiple equilibriums. Because the effects of the shocks are allowed to depend on the size and the sign of the shock, and also on the initial conditions, the impulse response functions are no longer linear, and it is possible to distinguish, for instance, between the effects of fiscal developments under different financial stress regimes.

Second, another advantage of the TVAR methodology is that the variable, by which different regimes are defined (s_t), can be an endogenous variable included in the VAR. Therefore, this makes it possible that regime switches may occur after the shock to each variable. In particular, the fiscal policy shock might either boost the output or increase the financial stress conditions that harm the prospects of economic growth, and the overall effect GDP of a fiscal expansion might became negative.

The threshold VAR can be specified as follows:

$$Y_t = A^1 Y_t + B^1(L) Y_{t-1} + \left(A^2 Y_t + B^2(L) Y_{t-1}\right) I\left[s_{t-d} > \gamma\right] + U_t \quad , \tag{1}$$

where Y_t is a vector of endogenous (stationary) variables, I is an indicator function that takes the value of 1 if, in our case, the financial stress s_t is higher than the threshold value γ, and 0 otherwise. The time lag d was set to 1. $B^1(L)$ and $B^2(L)$ are lag polynomial matrices, $A^1 Y_t$ and $A^2 Y_t$ represent the contemporaneous terms, because contemporaneous effects might also differ across the regimes. U_t are structural disturbances. We assume that the matrices A^1 and A^2 have a recursive structure.

We have used a recursive identification scheme for the VAR and included the following variables: GDP growth (y), inflation (π), the fiscal variable (f), the short-term interest rate (i), and the indicator for financial market conditions (s), for which we will use the Financial Stress Indicator (FSI) presented in section four. The VAR model in standard form is

$$Y_t = c + \sum_{i=1}^{p} V_i Y_{t-i} + \varepsilon_t \quad , \tag{2}$$

where Y_t denotes the (5×1) vector of the m endogenous variables given by $Y_t = \{y_t\ \pi_t\ f_t\ i_t\ s_t\}$, c is a (5×1) vector of intercept terms, V is the matrix of autoregressive coefficients of order (5×5), and the vector of random disturbances ε_t.

This particular ordering reflects some assumptions about the links in the economy. We order the FSI last which implies that the FSI reacts contemporaneously to all variables in the system. We assume that all new changes in both macroeconomic aggregates and economic policy that occur during one quarter are transmitted to financial markets within this quarter. The ordering of the fiscal variable after output is motivated by the need to identify the effects of automatic stabilizers in the economy. Hence, following Blanchard and Perotti (2002), we assume that all reactions of fiscal policy within each quarter (e.g. changes in government debt) are purely automatic because of implementation lags of fiscal policy measures. The interest rate shows up after the fiscal variable since the short-term interest rate can react contemporaneously to fiscal policy, but not vice versa.

The lag length of the endogeneous variables, p, is determined by the Schwarz information criteria, which attaches a larger penalty to the number of coefficients estimated in the model, hence we use only one or two lags given the low number of observations in the high stress regime. The main reason is that within the high financial stress regimes the number of observations is too low to allow estimating a VAR model with five variables and the conventionally used four lags.

We tested whether the threshold indicator is statistically significant. If the threshold values γ were known, the conventional F-test for the null hypothesis $A^2 = B^2(L) = 0$ would give reliable results. However, in our case the threshold value is not known a priori,[2] and the testing procedure involves non-standard inference, because γ is not identified under the null hypothesis of no threshold.

Therefore, we follow the procedure introduced by Hansen (1996). First, the TVAR model is estimated for all possible values of γ (to avoid over-fitting, the possible values were set so that at least 15% of the observations plus the number of coefficients is included in each regime), and the values of the Wald statistics testing the hypothesis of no difference between regimes are stored. Second, we constructed three test statistics: sup-Wald, which is the maximum value of the Wald statistics over all possible γ; the avg-Wald being an average of Wald statistics; and exp-Wald, which is the sum of exponential Wald statistics. To conduct inference, we simulated the empirical distribution of sup-Wald, avg-Wald and exp-Wald statistics with p-values obtained from 500 replications of the simulation procedure. The estimated thresholds were those that maximized the log determinant of the structural residuals U_t.

4.3.2 Nonlinear impulse responses

In a linear model, the impulse responses can be derived directly from the estimated coefficients and the estimated responses are symmetric both in terms of the sign and of the size of the structural shocks. Furthermore, these impulse responses are constant over time as the covariance structure does not change. However, these convenient properties do not hold within the class of nonlinear models as shown by Potter (1994) and Koop et al. (1996). The moving average representation of the TVAR is nonlinear in the structural disturbances U_t, because some shocks may lead to switches between regimes, and thus their Wold decomposition does not exist. Consequently, in contrast to linear models, we cannot construct the impulse responses as the paths the variables follow after an initial shock, assuming that no other shock hits the system. To cope with these issues, Koop et al. (1996) proposed nonlinear impulse response functions defined as the difference between the forecasted paths of variables with and without a shock to a variable of interest.

Formally, the nonlinear impulse responses functions (NIRF) are defined as

$$NIRF_y = (k, \varepsilon_t, \Omega_{t-1}) = E(Y_{t+k} | \varepsilon_t, \Omega_{t-1}) - E(Y_{t+k} | \varepsilon_t^{0,} \Omega_{t-1}) , \tag{3}$$

where Y_{t+k} is a vector of variables at horizon k, Ω_{t-1} is the information set available before the time of shock t. Note that the ε_t^0 denotes stochastic disturbance at time 0 that would occur under the no-shock scenario. This implies that there is no restriction regarding the symmetry of the shocks in terms of their sizes, because the effects of a ε_t shock depend on the magnitude of the current and subsequent shocks. Moreover, in the high-stress regime, the size of the fiscal shock matters, since a small shock is less likely to induce a change in the regime. Likewise, the impulse responses depend also on the entire history of the variables that affect the persistence of the different regimes.

Therefore, in order to get the complete information about the dynamics of the model, the impulse responses have to be simulated for various sizes and for the signs of the shocks. The algorithm proceeds as follows. First, the shocks for the periods from 0 to q are drawn from the residuals of the estimated VAR model. Then, for each initial value that is, for each point of our sample, this

2 Balakrishnan et al. (2009) suggest the value of one for the FSI to distinguishe the periods of high and low stress. Their judgement is based on the experience that such identification of stress periods mimics well the historical episodes of financial instability.

sequence of shocks is fed through the model to produce forecasts conditional on initial conditions. These steps are repeated for the same initial condition and the same set of residuals except for the shock to the variable of interest, which is set to +/-1 standard error and +/-2 standard errors at $t = 0$.

Second, we calculate the forecasts conditional on the shocks and on the initial conditions with and without an additional shock at $t=0$, and the difference between these two is the impulse response function. This procedure is replicated 500-times for each initial condition and the median, average and quantiles are saved. Then we compute averages over the initial conditions from each regime to get the impulse responses for both regimes.[3]

Because the number of observations in the high stress regime is rather low (ranging from 26 to 45), following Koop et al. (1996) we derive the confidence bands from the quantiles of the empirical distribution of the simulated impulse responses rather than assuming normality.

4.4 Variables and data

4.4.1 The dataset

A relevant issue with fiscal VARs is the choice of the variables that describe fiscal developments. For example, a discretionary increase in government revenues may have a different macroeconomic impact depending on which taxes are increased (labour versus consumption taxes), depending on whether a tax rate or the tax bases are modified, etc. At the same time, if one is data restricted, it is not possible to build VAR models with an excessive number of endogenous variables to describe fiscal policy.

We preferred to work with a parsimonious VAR structure to describe fiscal policy in the most aggregated form. Therefore, we used the government debt-to-GDP ratio because it reflects the developments both in government revenue and expenditure. Moreover, the government debt ratio captures also government actions that may not be fully reflected in the fiscal balance (e.g. purchase of financial assets, recapitalization of banking sector, the calling of previously issued government guarantees or any stock-flow adjustments) and has thus, in principle, a wider coverage of government actions than the fiscal balance. In addition, usually government debt is not a policy variable, with governments focussing more in the short run on the budget deficit rather than on government debt when forming their policies (e.g. governments typically announce budget deficit paths as their target).

The changes in the government debt ratio have an impact on the corporate sector expectations, consumption sentiment of households and on financial market conditions, since it provides information about not only the current fiscal policy but about past fiscal developments. In addition, the government debt ratio has a closer link to financial markets than the fiscal balance because it partly captures also the risk related to the refinancing of the outstanding stock of government debt, while influencing interest rates.[4]

The other variables that we include in the VAR are the already mentioned FSI (see next section for more details), GDP, the short-term interest rate and inflation. In some cases, and instead of the change in the debt ratio, we also used the budget balance ratio itself for robustness. However, on a

3 We used the WinRATS code provided by Nathan Balke, which we modified for our purposes.

4 The change in the debt ratio and the budget balance ratio are rather correlated for the countries in our analysis, see Figure A1 in Appendix.

quarterly basis such measure is more difficult to construct, for some countries than the debt ratio.

Regarding the time span we use a quarterly dataset, for the U.S., the U.K., Germany and Italy, for the period 1980:4-2009:4. Again, for some cases, instead of the FSI we also use alternative financial variables for the threshold in order to allow for a longer time span. The variables used in those cases were a measure of the stock returns and the so-called TED spread (the difference between the short-term interbank interest rate and treasury bills rate). For details see Appendix 1.

4.4.2 The Financial Stress Index

The FSI was developed by the IMF as an approximation to potential instability of financial markets (Cardarelli et al., 2010, substantially revised by Balakrishnan, 2009). The FSI contains three main components: (i) Bank related stress: beta of banking sector showing the perception of risk of the banking sector compared to other sectors in the economy, the TED spread (difference between the short-term interbank interest rate and treasury bills rate) and the inverted term structure. (ii) Securities related stress: corporate bond spread, stock market returns and stock-market volatility. (iii) Exchange rate stress: exchange rate volatility. The FSI index is then constructed as a sum of normalized values of all these sub-components.[5] The larger value of the FSI, the higher is the stress during each period. The authors have shown, that these components are relatively uncorrelated and, importantly, adding different variable does not change the resulting path of the FSI significantly.[6]

Cardarelli et al. (2010) describe the effects of the FSI and its sub-components on output. Based on their findings the most important effects on output occur in the periods of financial stress connected with the banking sector. Baxa, Horvath and Vasicek (2010) studied the reaction of central bank inflation targeting to financial stress using the augmented Taylor rule with time-varying coefficients. They found that these central banks normally do not react to financial stress, however, their behaviour changes in times of large and longer stress such as the Bank of England during the ERM crisis and the 2008-2009 crisis, for example.

4.4.3 Fiscal developments' overview

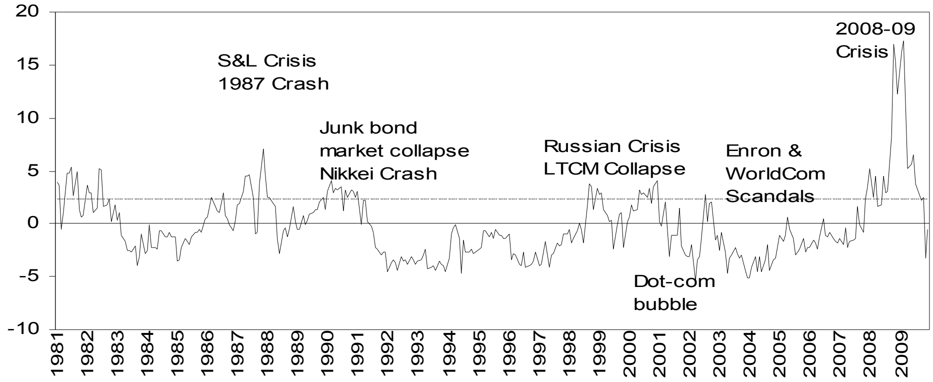
Figure 2 provides some evidence about fiscal policies in the U.S., the U.K., Germany and Italy in the period 1970-2009, based on the annual national accounts data from the European Commission Ameco database. In order to capture the main fiscal developments during this period we plot two charts: the first one with the general government debt-to-GDP ratio on the left-hand side axis and with government revenue and expenditure ratios on the right-hand side axis; the second one with the general government balance on the left-hand side axis and government debt on the right-hand side axis.

5 The version of the FSI in Cardarelli et al. (2010) is constructed by taking the average of the components after adjusting for the sample mean and standardizing by the sample standard deviation. Then, the index is rebased so that it ranges from 0 to 100. Episodes of high stress are identified as those periods when the FSI is more than one standard deviation above its trend determined by the Hodrick-Prescott filter.

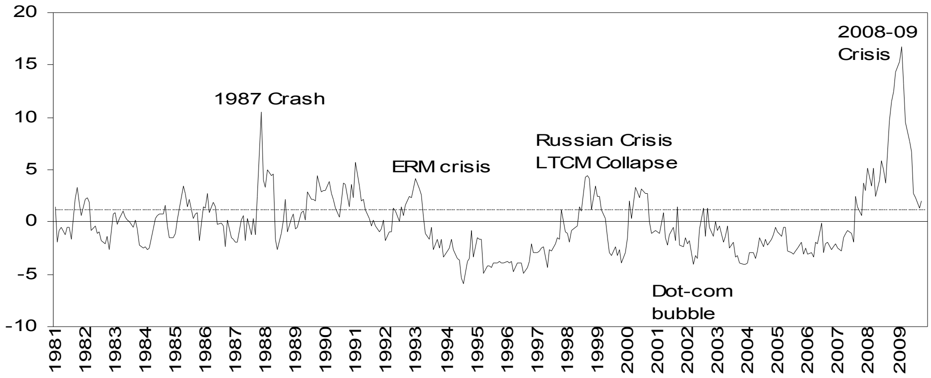
6 Regarding the exchange rate component we do not observe, for Germany and Italy, significant changes around the adoption of the euro in 1999. For Italy, some volatility can be seen after the exiting of the Italian Lira and of the British Pound from the European Exchange Rate Mechanism on September 1992.

Figure 1 – Financial Stress Indicator

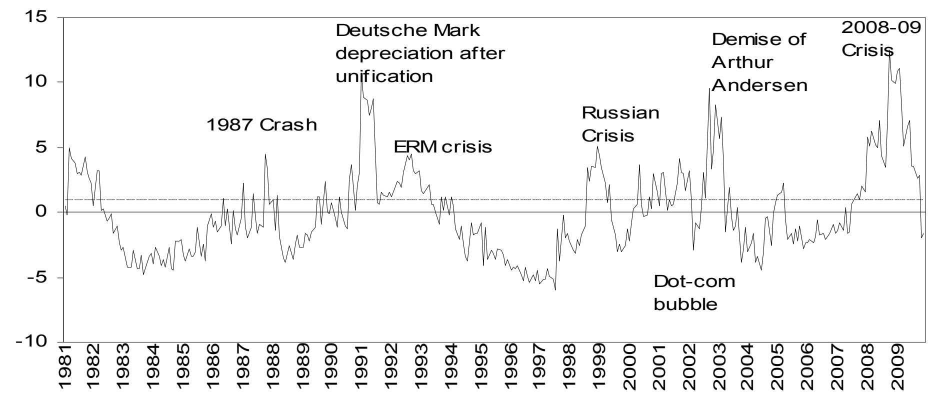
1.1. U.S.



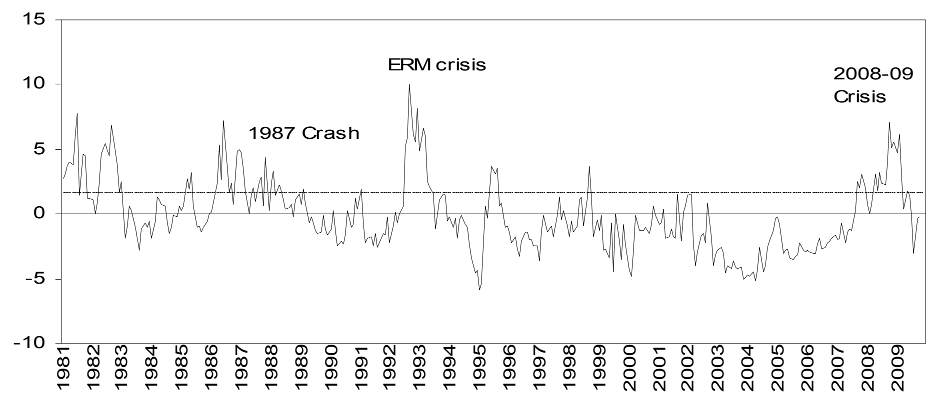
1.2. U.K.



1.3. Germany



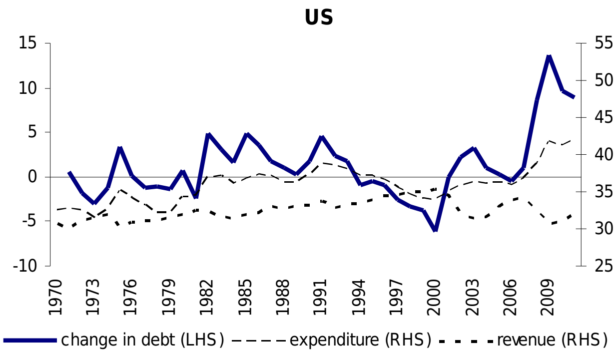
1.4. Italy



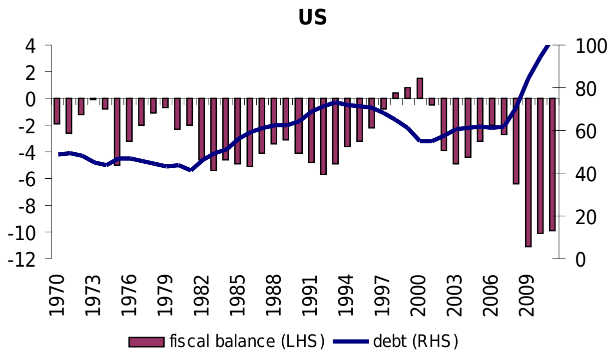
Source: Cardarelli et al. (2009) and the authors.

Figure 2 – General government debt, revenue, expenditure
and fiscal balance developments, in % of GDP

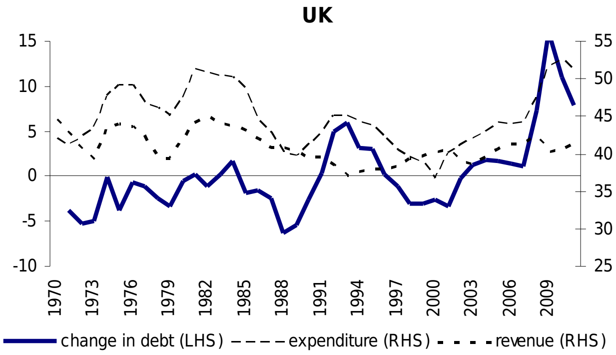
2.1. US – spending and revenue



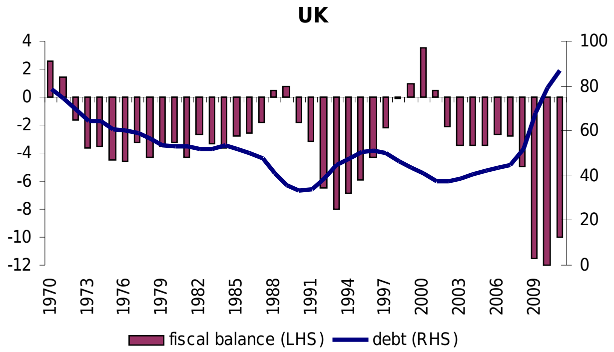
2.2. US – debt and deficit



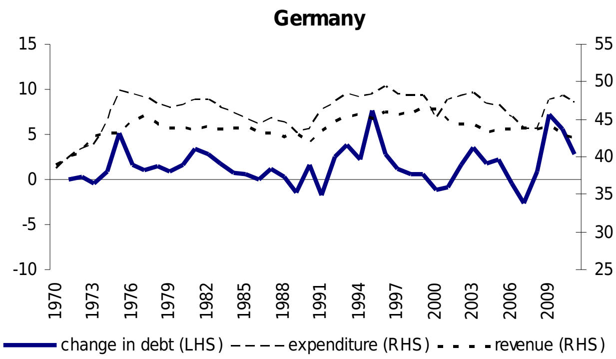
2.3. UK – spending and revenue



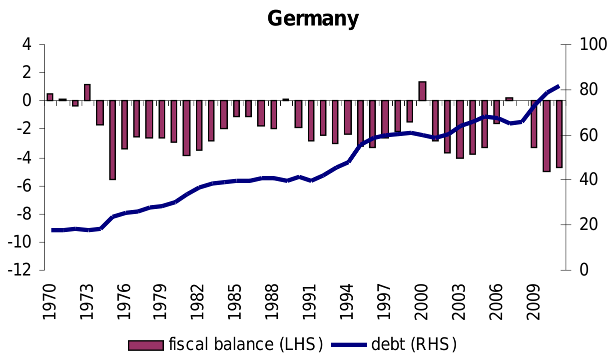
2.4. UK – debt and deficit



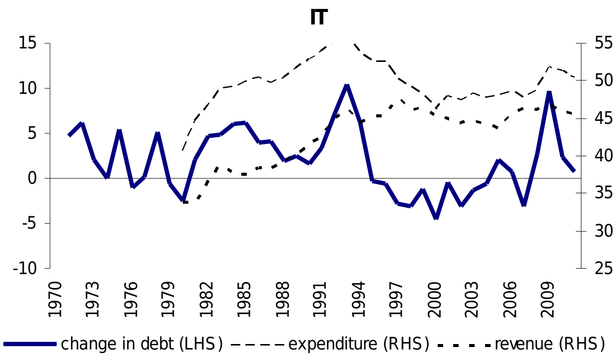
2.5. Germany – spending and revenue



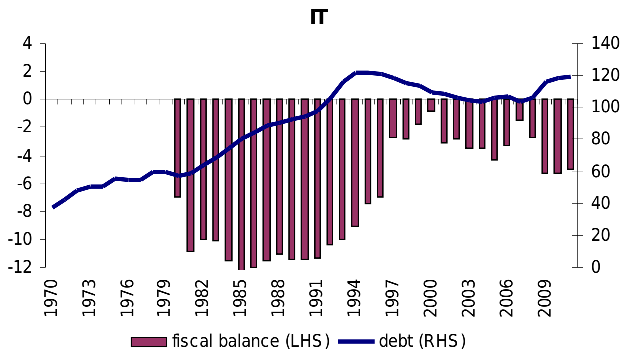
2.6. Germany – debt and deficit



2.7. Italy – spending and revenue



2.8. Italy – debt and deficit



Source: European Commission Ameco.

In the U.S., the periods with high financial stress broadly correspond to recessions. This is the case in particular for the recessions identified by the NBER between 1981Q3-1982Q4, 1990Q3 – 1991Q1, 2001Q1 – 2001Q4 and the latest recession that started in 2007Q4. However, the financial stress was identified also in the non-recession periods in 1987Q3, 1988Q1 and 1999Q2. The stress in the financial markets in 1987Q3 is related to the event "Black Monday", 19 October 1987, when the stock market in Hong Kong crashed and the effects spread globally. The second non-recession period of tension in 1988Q1 could be linked to the savings and loan crisis in the US. In that year, several banks located mainly in Texas and California went under (e.g. First Republic Bank, American Savings and Loan Bank and First City National Bank).

The government debt ratio was gradually declining until 1981 when a recession hit the U.S. economy and the debt ratio started to increase. In August 1982, the Congress approved the Tax Equity and Fiscal Responsibility Act and the previous tax cuts, which were implemented in the Economic Recovery Tax Act in 1981, were heavily reversed. The recession finished in the autumn of 1982, but the debt ratio continued to increase until 1990 when another recession occurred. In the autumn of 1990, the U.S. government enacted legislation which targeted a cumulative deficit reduction of about $500 billion over five years. In addition, the government improved also the fiscal framework and prepared the Budget Enforcement Act, which introduced new fiscal rules to limit future budget deficits and discretionary expenditures. The recession finished in the spring of 1991 and the debt-to-GDP ratio peaked two years after, in 1993, at about 72%.

The following recovery brought the debt ratio on a declining path that lasted until 2001 when a recession emerged and contributed to the ensuing fiscal deterioration. Despite the fact that this recession was over already by the end of the same year, the government debt ratio gradually increase to 62% of GDP in 2007, when the subprime debt crisis severely affected the U.S. economy. In 2008, the U.S. administration faced a serious recession and adopted a fiscal stimulus package consisting of federal tax cuts and spending increases of about 5% of GDP. As a consequence, the general government deficit jumped to about 11% of GDP in 2009, the highest number since 1970 and well above the deficits of 5.4% of GDP in 1983 and 5.7% of GDP in 1992, which can be linked to previous recessions.

Interestingly, for the U.S. Favero and Giavazzi (2007) point to different effects of exogenous tax policy shocks on output in the period 1980-2006, when compared to the previous period. In the 1950s, 1960s and 1970s the contractionary effect of a tax hike was larger when monetary policy shocks, government spending, and oil prices were endogeneized in a model that included the level of the debt and the government intertemporal budget constraint. Since the beginning of the 1980s, when the burden of debt stabilization falls on expenditure, an exogenous increase in taxes was compensated by a subsequent expenditure accommodation. This could explain why, analyzing the effects of shocks in a model with endogenous monetary policy, government spending, oil prices, and fiscal policy, produced much smaller output effects. Favero and Giavazzi (2007) argued that in fact since the beginning of the 1980s, an initial positive tax shock is accompanied by further tax changes in the opposite direction in the U.S. Following the initial shock taxes decline and the effect on the budget is compensated by increases in spending.

In the U.K., government debt had been continuously declining from high levels of around 80% at the beginning of the 1970s to around 33% in 1990. A particularly strong fiscal consolidation was carried out in 1988 and 1989 when the fiscal balance recorded surpluses of about 0.5 and 0.8 % of GDP, respectively.

However, the orientation of British fiscal policy has changed several times since the 1970s. In the 1970s, fiscal policy was the key policy instrument used for aggregate demand management. When a new conservative government took office in 1979 keynesianism was replaced by monetarism as the leading economic paradigm. The fiscal policy strategy changed and focused on reducing the size of the government in the economy in addition to suppressing the role of fiscal policy in demand management. In the early 1970s, an oil price shock helped in causing larger contractions in output and led to stagflation. The U.K. economy contracted around 2% in the beginning of 1974. On the contrary, the recession was much bigger in the beginning of 1980 when GDP dropped by almost 5%. The 1980 recession was deep also when compared to the recession in the 1990 and 1991 when the economy contracted by less 2 %.

In Germany it is possible to identify a few periods of fiscal consolidation episodes, notably the period 1982-83 when the cyclically adjusted primary budget balance improved more significantly (see also Figure 2 for the overall fiscal balance). The debt ratio increased gradually from a very low level, less than 20% of GDP in 1970, to about 70% of GDP over the sample period with only four relatively short periods of debt ratio reduction in 1979, around 1989-1991, in 2000-2001 and 2006-2007 which coincide with the peaks of the business cycle. In 1979, the real GDP growth rate reached almost 5% and in 1990 peaked at 5.25% in West Germany. However, the period which followed the German reunification in 1990, in which the exchange rate stress component of FSI was particularly high, must be interpreted with caution, because the German economy had to cope with the economic transition of the former East Germany from planned to market economy. The economic transition required large amounts of public spending which stimulated an economic boom in several German regions. The following peaks of real GDP growth rate that led to GDP ratio reductions were recorded in 2000 and 2006 when the growth rate reached 3.2%. From a fiscal policy perspective, important changes followed the ambitious and large tax reform in 2000 in which the German government passed the most ambitious tax reform and the tax burden was reduced for both individuals and companies. As a consequence, the revenue-to-GDP ratio decline by almost 3 p.p. of GDP between 1999 and 2008. The changes in the German fiscal policies are more complex due to fiscal federalism, where fiscal decisions of local governments play a more important part.

In Italy, the debt ratio increased from about 37% of GDP in 1970 to about 122% of GDP in 1994, then declined to about 104% of GDP by 2004 and further increased to 115% of GDP in 2009. This was mainly due to a more relaxed fiscal policy in the 1980s with the occurrence of budget deficits of 10-12% of GDP each year. The consolidation effort started to materialize in 1995 when the debt ratio declined by 0.3% of GDP. One of the main drivers of the Italian fiscal consolidation in the 1990s was the effort to fulfil the Maastricht fiscal criteria, which are necessary to qualify for the euro area membership. For more details on fiscal consolidation process that was characterised by a large number of corrective measures with only temporary effects, see, for instance, Balassone et al. (2002).

The period of fiscal prudence between 1995 and 2004 delivered a notable reduction in the government debt ratio, which declined by about 18 p.p. during that period. This reduction of government debt decreased, inter alia, government interest expenditures from typically around 11-12% of GDP in the 1980s to less than 5% of GDP since 2004. The interest payments usually constituted a substantial part of government expenditures in the past. For example in the 1980s, the interest expenditures corresponded to about 70% of the overall fiscal deficit and in the beginning of the 1990s, the ratio of government interest expenditures to GDP typically exceeded the fiscal deficit ratio, allowing the delivery of primary budget surpluses. In those years, the financing of government interest expenditures consumed about 1/3 of total government revenues.

Contrary to the German experience, where the debt reduction occurs in a short two-year period that reflect mostly the business cycle, the debt reduction in Italy has a different pattern mainly due to the downward trend in nominal interest rates and consolidation efforts in mid-1990s (see Figure 1). A similar patter can be found in the UK, where the debt ratio declined in almost twenty consecutive years since 1970 with only one interruption of this declining trend in 1984. While economic growth seems to be the major factor of debt reductions in Germany, the decline of interest expenditures also played a significant role in the Italian fiscal consolidation efforts. For an assessment of fiscal consolidation episodes in the EU countries, see Afonso (2010).

4.5 Results

4.5.1 Testing the Threshold VAR model

We tested whether the data indicate the presence of a statistically significant threshold γ as defined by the values of the financial stress index, and whether the optimal threshold values are reasonable in terms of identifying high and low stress periods that will be related to output fluctuations. Because the observed persistence of the financial stress index is very low, when comparing to fluctuations of other macroeconomic variables, reasonable values of the threshold would have lead to a segmentation of periods with high financial stress. Therefore, we have determined the threshold from the FSI smoothed by a 3-period moving average.

Our estimated threshold values range from 0.92 for Germany to 2.38 for the U.S. and the threshold is significant with a p-value often less than 0.0001 for all the Wald statistics (Table 1).[7]

The threshold splits the sample into a high stress regime with about one fourth of observations (from 24 to 39) and a low stress regime with the remaining portion. Such division seems to be well in line with the fact that the duration of expansions is higher than the duration of recessions. The number of observations of the high stress regime makes the VAR model less parsimonious in this regime. To address possible biases in our results, caused by the limited number of observations within the high stress regime, we estimated the threshold VAR also for other variables representing instability in financial markets, whose time series went further back in time than the FSI (available since the fourth quarter of 1980). These variables were: a measure of stock returns, and the TED spread measuring the spread between the interest rate on Eurodollar papers and treasury bills and for the U.S. also the spread between the commercial paper rate and the treasury bills. Such experiments confirmed our main findings about the effects of fiscal policy in both regimes.

Therefore, in Table 1 we report the estimated thresholds for each country, both using the FSI indicator and alternative financially related variables.

We can see in Table 2 that high stress periods identified using the estimated thresholds are more frequent than recessions. However, all recessions in all countries have their counterpart among the high financial stress periods. Additionally, the average annual output growth rates are lower in high stress periods than in low stress periods.

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7 The optimal values maximize the log determinant of residuals for all countries except the U.K., where the value maximizing the Wald statistics was chosen. In this case, the maximized log determinant of residuals implied a threshold equal to 0.2585, but the maximum Wald statistics was for the threshold γ = 1.2369. The latter value is more in line with other countries and also gives a similar share of observations in both regimes.

Table 1 – Thresholds per country

Threshold variable	Estimated Threshold	Sup-Wald	Avg-Wald	Exp-Wald	VAR order	Sample	N. observations Low Stress	High Stress
				U.S.				
FSI	2.3822	100.85	62.11	47.11	1	1980Q4 - 2009Q4	88	24
TED	1.62	331.49	102.93	161.32	2	1971Q1 - 2009Q4	125	35
Stock Returns	-0.1622	166.64	138.97	78.54	2	1956Q1 - 2009Q4	167	47
				U.K.				
FSI	1.2369	179.65	109.37	85.81	1	1980Q4 - 2009Q3	81	29
TED	0.3143	200.86	132.12	96.58	1	1979Q1 - 2009Q3	92	29
Stock Returns	1.2531	138.23	111.17	65.77	2	1978Q2 - 2009Q3	77	44
				Germany				
FSI	0.9167	121.63	94.81	57.75	2	1980Q4 - 2009Q4	77	39
Stock Returns	1.3067	148.51	105.04	70.27	2	1979Q1 - 2009Q4	79	72
				Italy				
FSI	1.725	72.51 (0.016)	47.2 (0.136)	32.8 (0.016)	1	1980Q4 - 2009Q3	113	26
TED	-0.4898	114.73	90.72	53.94	1	1979Q1 - 2009Q3	87	35

Notes: TED - spread between the short-term interbank interest rate and the treasury bills rate. Stock returns: US – based on Dow Jones Industrial Index; UK – based on the Financial Times Stock Exchange (FTSE) 100 index; Germany - based on the IMF IFS share prices indicator. p-values were always less than 0.0001, if not, their values are in parentheses.

Table 2 – High financial stress and recessions

U.S.		***NBER recessions***		**Annual output growth rates**		
High Stress[1]		Peak	Trough		High Stress	Low Stress
1981Q3	1983Q1	1981Q3	1982Q4	***average***	***0.33***	***3.50***
1987Q3				min	-3.90	-0.27
1988Q1				max	4.81	8.15
1990Q3	1991Q2	1990Q3	1991Q1	stdev	2.72	1.43
1999Q2						
2000Q4	2001Q2	2001Q1	2001Q4			
2008Q2	2009Q4	2007Q4	2009Q2			
U.K.		***ECRI Recessions***				
High Stress[1]		Peak	Trough		High Stress	Low Stress
1982Q2				***average***	***1.68***	***2.77***
1986Q1				min	-4.69	-2.55
1988Q2	1988Q4			max	5.29	4.94
1989Q2	1991Q3	1990Q2	1992Q1	stdev	2.39	1.38
1993Q1	1993Q3					
1998Q4	1999Q2					
2000Q4	2001Q1					
2008Q2	2009Q3	2008Q2				
Germany		***ECRI Recessions***				
High Stress[1]		Peak	Trough		High Stress	Low Stress
1981Q3	1982Q3	1980Q1	1982Q4	***average***	***0.92***	***2.12***
1991Q1	1993Q4	1991Q1	1994Q2	min	-5.01	-0.69
1999Q1	1999Q3			max	6.12	5.44
2001Q1	2002Q2	2001Q1	2003Q3	stdev	2.38	1.26
2002Q4	2003Q4					
2008Q2	2009Q4	2008Q2	2009Q1			
Italy		***ECRI Recessions***				
High Stress[1]		Peak	Trough		High Stress	Low Stress
1981Q3	1983Q2	1980Q2	1983Q2	***average***	***-0.14***	***1.49***
1986Q3	1988Q1			min	-6.63	-0.37
1992Q4	1993Q4	1992Q1	1993Q4	max	4.32	4.59
2008Q2	2009Q3	2007Q3		stdev	2.83	1.32
Per memory						
Euro area		***CEPR Recessions***				
		Peak	Trough			
		1974Q3	1975Q1			
		1980Q1	1982Q3			
		1992Q1	1993Q3			
		2008Q1	2009Q2			

Sources: NBER, US Business Cycle Expansions and Contractions, http://www.nber.org/cycles/cyclesmain.html
CEPR, Euro Area Business Cycle Dating Committee, http://www.cepr.org/data/dating/.
Economic Cycle Research Institute (ECRI) www.businesscycle.com.
[1] identified using the estimated thresholds.

4.5.2 The effects of fiscal shocks

Figure 3 reports the median impulse response functions of a fiscal policy shock, both for a high and for a low financial stress regime. We opted for the median impulse response functions and the respective confidence bands derived from the empirical distribution of the responses rather than from the normal distribution due to the lower number of observations namely in the higher stress regime sample.[8] Broadly, the responses of output growth to a fiscal shock are positive in both regimes and in all countries in our sample, although in some cases the response is either initially negative or uncertain within the first few quarters after the shock.

In the U.S. the responses of output growth to an increase in the budget deficit are similar in terms of their peak effect in both regimes. However in the high stress regime, the impulse response is negative in the first quarter after the shock. On the other hand, the increase in output growth is faster in comparison to the low stress regime. The impulse response functions are significant at 50%, in the high stress regime after four quarters. When the budget balance is used instead of the change of the debt ratio in the threshold VAR, the results change only slightly. The response of output to a fiscal shock does not have the initial small negative effect and it is always positive. The low stress regime is different and the effect of a positive fiscal shock is temporarily negative and it turns into a positive effect after three quarters. In this specification the different behaviour is caused by a switch in contemporaneous terms of the VAR: the FSI drops after a positive fiscal shock in the high stress regime, but it temporarily increases in the low stress regime. Thus, our result of stronger fiscal effects in high financial stress survives this sensitivity check.[9]

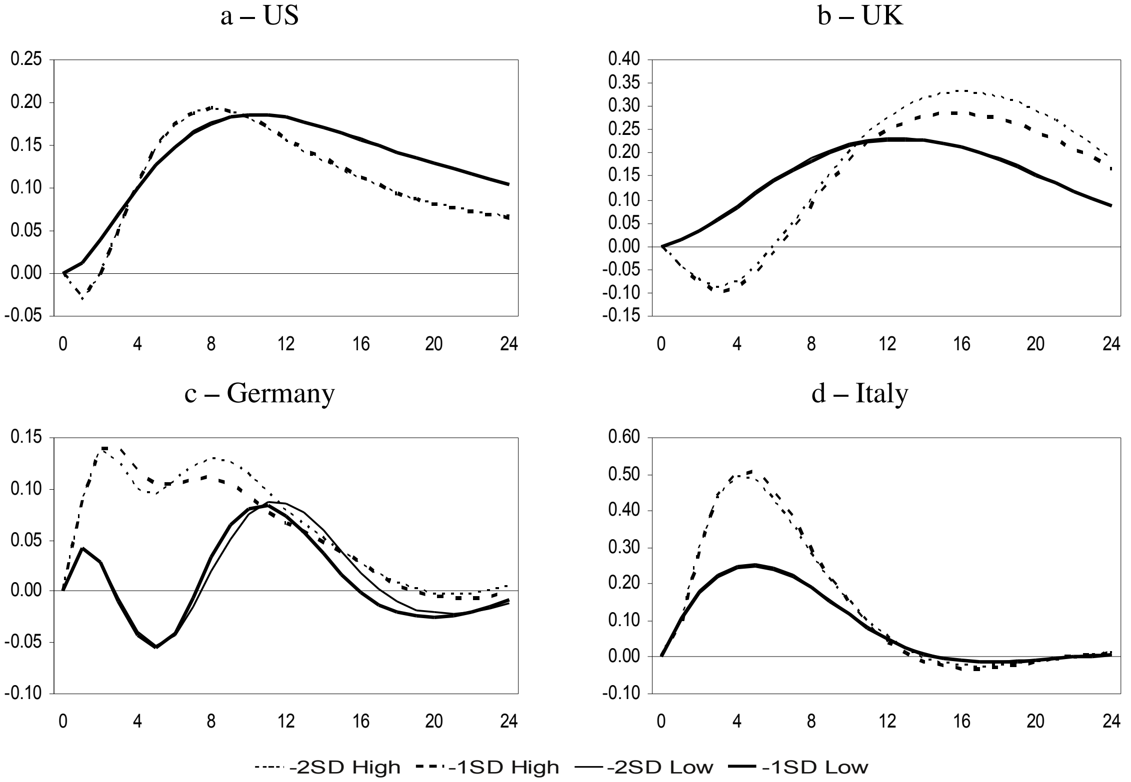
In the U.K., fiscal policy causes an increase in output growth when the economy is in the low stress regime. However, the impulse response of output in the high stress regime exhibits a similar pattern to the U.S., and initially output growth decreases. This decrease lasts for six quarters and the 75% quantile of the simulated impulse responses is even below zero (see Appendix 3, Figure A3.1). Contrary to the low stress regime, financial stress does not decrease in response to a fiscal shock in the high stress regime.

8 To assess uncertainty, we report the 25%-75% and 5%-95% quantiles along with the median impulse response function used in the main text in Appendix 3 (Figure A3.1). Both point estimates and the quantiles evolve over time, which implies that even though some impulse responses are insignificant at the 90% level, they still might be significant at the same level within several periods. Hence their interpretation is not so straightforward as in time-invariant VAR models.

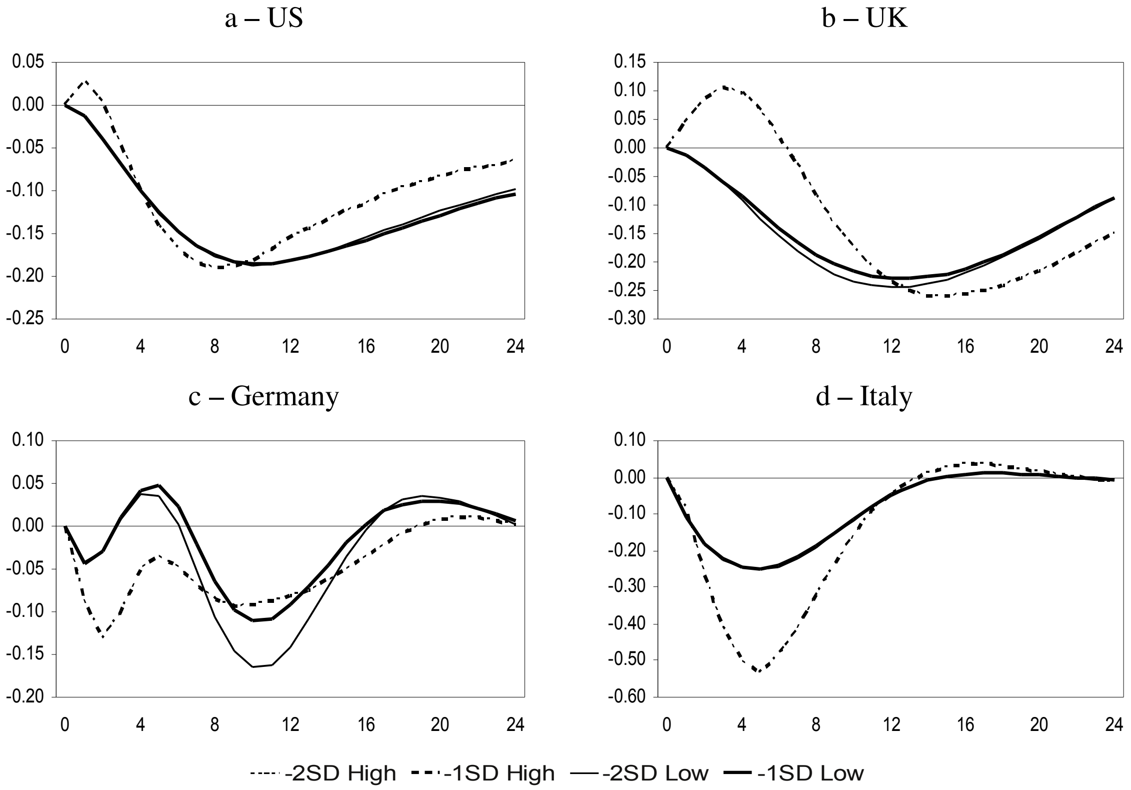
9 For robustness, we also replaced the change in the debt ratio by a change in the debt itself and the results were unchenged. We also used the first differences of GDP and of the price level together with the budget balance. This was the only specification when a 1% fiscal shock had larger effects on output growth in a higher stress regime than in the lower one. Again, the amplitude of the impulse response of output was reached earlier in the high stress regime than in the lower one. Furthermore we estimated the effects of a very large shock of 5 standard deviation (SD) in the high stress regime, corresponding to about 3.5% of GDP. The magnitude of the effect was roughly proportional to the 2 SD shock, but the peak was reached even faster, within 6 quarters after the shock (for a 2 SD shock it was 8 quarters).

Figure 3 - Fiscal Shock, Response of Output Growth

3.1 – Positive fiscal shock



3.2 – Negative fiscal shock



Note: Fiscal shocks rescaled to initial +/-1%GDP shock

For Germany, the effect of a positive fiscal policy shock on output is positive, when the economy is in the high stress regime. The response of output in the low stress regime is oscillating during the first eight quarters from a positive to a negative impact, but then the response becomes positive. However the response in the low stress regime is insignificant, although the distribution of impulse responses indicates that it is more likely positive. Table 3.1 reports multipliers confirming that fiscal policy has larger effects on output in the high stress regime than in the low stress regime. The different responses in both regimes are caused by a number of factors. First, the dynamics of the fiscal shock is different and somewhat increasing endogenously after the initial shock in the high stress regime, and monotonously decreasing in the low stress regime. Second, the financial stress indicator reacts differently. When the economy is in a high financial stress regime, it increases to a value above 1, and it is positive for the first three periods and negative afterwards. This explains the temporary decrease in the response of output growth. In the low stress regime the financial stress indicator decreases in a hump-shaped pattern.[10]

The results for Italy show that notwithstanding the high level of government debt, the responses of output to a fiscal shock follow the Keynesian pattern. In both regimes, the response of output is positive with a hump-shaped pattern.

As far as the effects of the size of the fiscal shock are concerned, both Figures 3.1 and 3.2 do not provide evidence of important asymmetries between small and large shocks with the exception of Germany. Moreover, one and two standard deviations shocks practically coincide in the U.S. and in Italy.

When a negative fiscal shock is considered, responses coincide in the high stress regime and only relatively smaller differences arise in the low stress regime. Germany is somewhat different. The impulse responses of positive fiscal shocks are slightly dissimilar, but in terms of the cumulative multipliers over three years the differences are negligible. However, large fiscal contractions in the low stress regime lead to non-proportionally larger effects on output and their cumulative multipliers are almost twice the ones corresponding to small fiscal shocks (see also Table 3.1).

We did a couple of robustness checks. First, we replaced the change in the debt ratio by a change in the debt itself and the results were unchanged. Similarly, when the budget balance is used instead of the change in the debt ratio, the results change only slightly in all four countries.

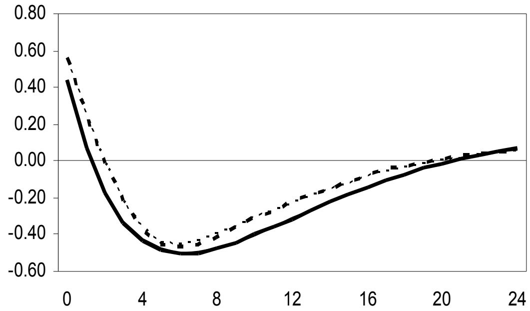
The responses of financial stress to fiscal shocks are presented in Figure 4. The results show that fiscal policy shocks decrease the financial stress, when the economy is initially in the high stress regime, although there is an initial increase of financial stress within the first few quarters. Only the U.K. is somewhat different, and the impulse response of the FSI does not decrease to negative values but, on the contrary, it implies increased financial stress both in the short and in the long term. This long-term increase of FSI is larger for large positive fiscal shocks, suggesting that fiscal policy does not decrease financial stress in the U.K. In the low stress regime the response of financial stress is negative with a hump-shaped pattern in all countries, except for Italy, where the financial stress indicator reacts only moderately.

10 For example, a change in the debt ratio could reflect efforts to reduce financial stress, rather than to stabilise economic growth.

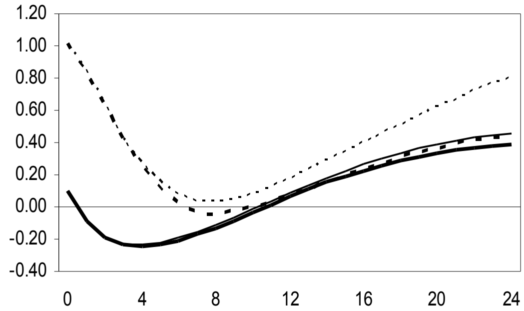
Figure 4 - Fiscal Shock, Response of FSI

4.1 – Positive fiscal shock

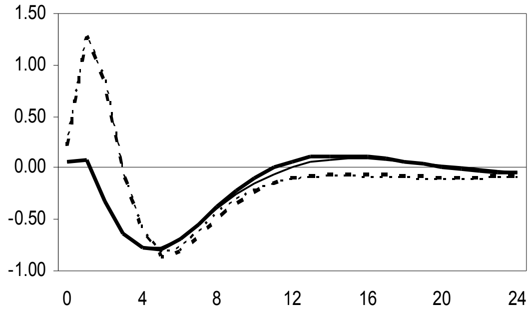
a – US



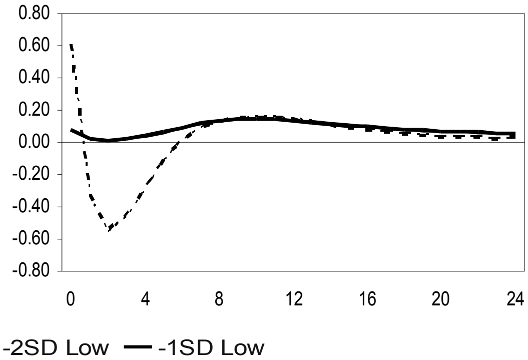
b – UK



c – Germany

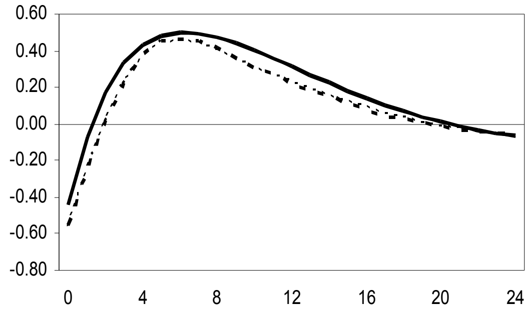


d – Italy

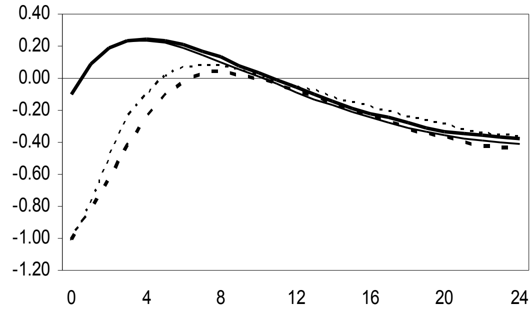


4.2 - Negative fiscal shock

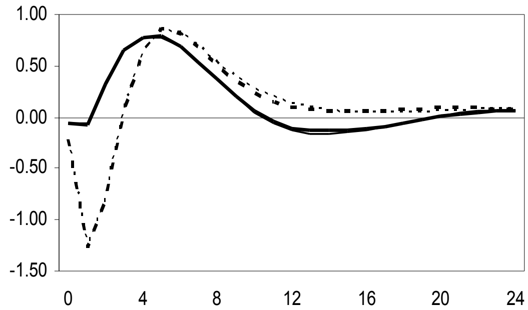
a – US



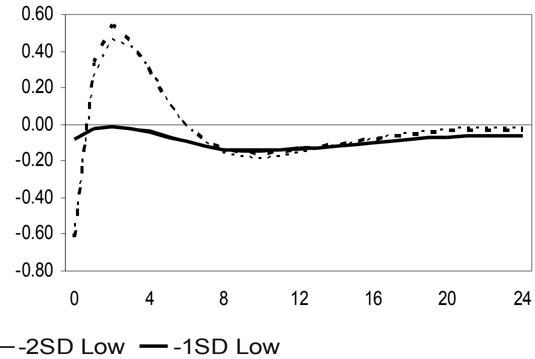
b – UK



c – Germany



d – Italy



Note: Fiscal shocks rescaled to initial +/-1%GDP shock

Some additional points are worthwhile mentioning. First, a positive fiscal shock leads to a temporary increase of financial stress in Germany, but after few quarters the path of FSI reverts and follows the scenario related to the low stress regime. Second, the financial stress indicator reacts only moderately in Italy, when the economy is in the low stress regime.

Tables 3.1 and 3.2 summarize the values of the multipliers for the responses of output and FSI at one, two and three years after a fiscal shock, and also a cumulated response over three years. The impulse responses are normalized to the same size of the initial fiscal shock set to 1% of GDP for a direct comparison between two (High and Low stress) regimes and different signs and sizes. We use 1SD and 2SD as proxies for small and large shocks.

The size of fiscal multipliers varies across countries and across regimes. The multipliers are largest in Italy with a size of the cumulative multiplier after three years of about 0.82-0.87 for the high stress regime and 0.48-0.49 for the low stress regime. In Germany the cumulative fiscal multiplier is 0.3 in the high stress regime and almost zero when the economy is initially in the low stress regime, implying strong crowding-out effects in the economy. For the U.S. the cumulative multipliers are between 0.45-0.46 with minor differences between signs and sizes of shock. The U.K. has the lowest effects of a fiscal policy shock on output growth in the high stress regime, with the cumulative multiplier over three years being between 0.22 and 0.3. Interestingly, if the fiscal shocks occur in the low financial stress regime, the cumulative multipliers are around 0.50-0.54.

Table 3.1 – Responses of output to a 1% of GDP fiscal shock

	4 Quarters		8 Quarters		12 Quarters		Cumulative (12 quarters)	
	2SD	1SD	2SD	1SD	2SD	1SD	2SD	1SD
U.S.								
Positive Shock								
High	0.103	0.105	0.193	0.194	0.153	0.157	0.449	0.456
Low	0.100	0.100	0.177	0.176	0.182	0.182	0.460	0.458
Negative Shock								
High	-0.101	-0.101	-0.190	-0.190	-0.155	-0.155	-0.445	-0.445
Low	-0.100	-0.100	-0.177	-0.175	-0.182	-0.182	-0.459	-0.457
U.K.								
Positive Shock								
High	-0.076	-0.091	0.103	0.087	0.275	0.246	0.301	0.242
Low	0.088	0.085	0.189	0.184	0.230	0.229	0.507	0.497
Negative Shock								
High	0.097	0.097	-0.085	-0.085	-0.234	-0.234	-0.222	-0.222
Low	-0.091	-0.085	-0.203	-0.186	-0.243	-0.229	-0.537	-0.500
Germany								
Positive Shock								
High	0.099	0.117	0.130	0.113	0.079	0.067	0.308	0.296
Low	-0.039	-0.042	0.020	0.033	0.086	0.074	0.068	0.065
Negative Shock								
High	-0.051	-0.051	-0.085	-0.085	-0.082	-0.082	-0.218	-0.218
Low	0.037	0.041	-0.107	-0.065	-0.141	-0.091	-0.211	-0.115
Italy								
Positive Shock								
High	0.491	0.498	0.281	0.295	0.054	0.043	0.826	0.836
Low	0.248	0.244	0.193	0.190	0.053	0.049	0.494	0.483
Negative Shock								
High	-0.501	-0.501	-0.324	-0.324	-0.045	-0.045	-0.871	-0.871
Low	-0.247	-0.244	-0.192	-0.187	-0.049	-0.048	-0.488	-0.479

Table 3.2 – Responses of financial stress to a 1% of GDP fiscal shock

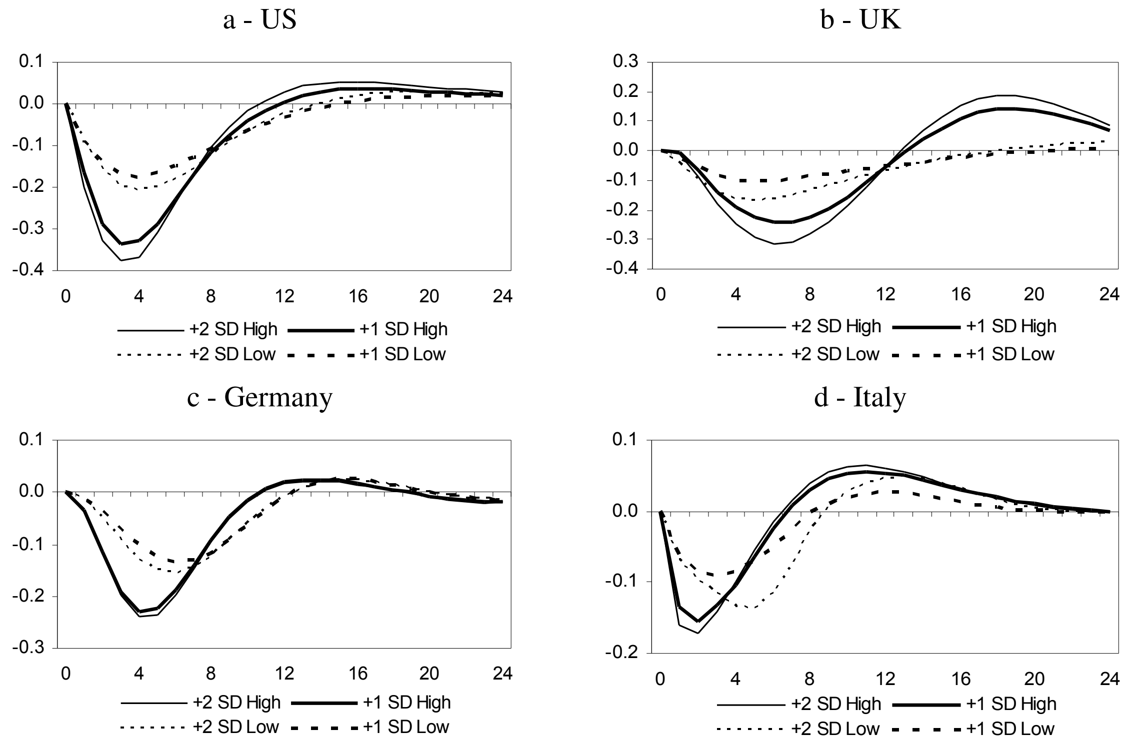
	4 Quarters		8 Quarters		12 Quarters		Cumulative (12 quarters)	
	2SD	1SD	2SD	1SD	2SD	1SD	2SD	1SD
				U.S.				
				Positive Shock				
High	-0.371	-0.377	-0.399	-0.411	-0.228	-0.224	-0.998	-1.013
Low	-0.428	-0.430	-0.475	-0.476	-0.313	-0.313	-1.215	-1.219
				Negative Shock				
High	0.376	0.377	0.403	0.411	0.232	0.224	1.010	1.013
Low	0.426	0.430	0.470	0.476	0.308	0.314	1.204	1.220
				U.K.				
				Positive Shock				
High	0.264	0.243	0.030	-0.039	0.179	0.073	0.473	0.276
Low	-0.237	-0.245	-0.108	-0.129	0.086	0.062	-0.259	-0.311
				Negative Shock				
High	-0.094	-0.243	0.076	0.039	-0.054	-0.073	-0.072	-0.276
Low	0.236	0.246	0.100	0.129	-0.085	-0.056	0.251	0.319
				Germany				
				Positive Shock				
High	-0.622	-0.642	-0.448	-0.498	-0.110	-0.104	-1.180	-1.244
Low	-0.781	-0.780	-0.397	-0.382	0.014	0.067	-1.164	-1.095
				Negative Shock				
High	0.635	0.642	0.533	0.498	0.138	0.104	1.306	1.244
Low	0.778	0.779	0.367	0.371	-0.129	-0.103	1.015	1.047
				Italy				
				Positive Shock				
High	-0.288	-0.287	0.126	0.131	0.147	0.140	-0.014	-0.016
Low	0.032	0.041	0.133	0.136	0.138	0.134	0.303	0.310
				Negative Shock				
High	0.303	0.287	-0.155	-0.131	-0.160	-0.140	-0.013	0.016
Low	-0.036	-0.044	-0.137	-0.136	-0.136	-0.134	-0.309	-0.313

4.5.3 *The effects of financial stress shocks*

The responses to a shock in the financial stress indicator are in accordance with our expectations. The effect on output is negative and it erodes after 6-10 periods, when it temporarily becomes positive, namely in the U.K. and in Italy. As we can see in Figure 5, there is some evidence of asymmetric reactions between large and small shocks, and also between the two regimes.

Figure 5 – Responses of Output Growth to a Positive Shock in Financial Stress

High and Low Regimes



Note: The impulse responses were rescaled to the size of the shock of one unit of FSI.

Table 4 reports the values of the impulse responses of output at different horizons. Several conclusions can be drawn from these results. First, the effect on output growth of increased financial stress tends to be larger in the high stress regime than in the low stress regime, especially at the horizon of 8 quarters after the shock. In the high stress regime, the output falls more than proportionally after a +2SD shock in comparison to +1SD shock in the U.S. and in the U.K., whereas in Germany and in Italy the impact of a financial stress shock is, in principle, proportional to the size of the initial shock. The responses to negative financial stress shocks are in principal proportional to a +1SD shock at both horizons, except in the U.K., where the impact of -2SD shock is less than proportional.

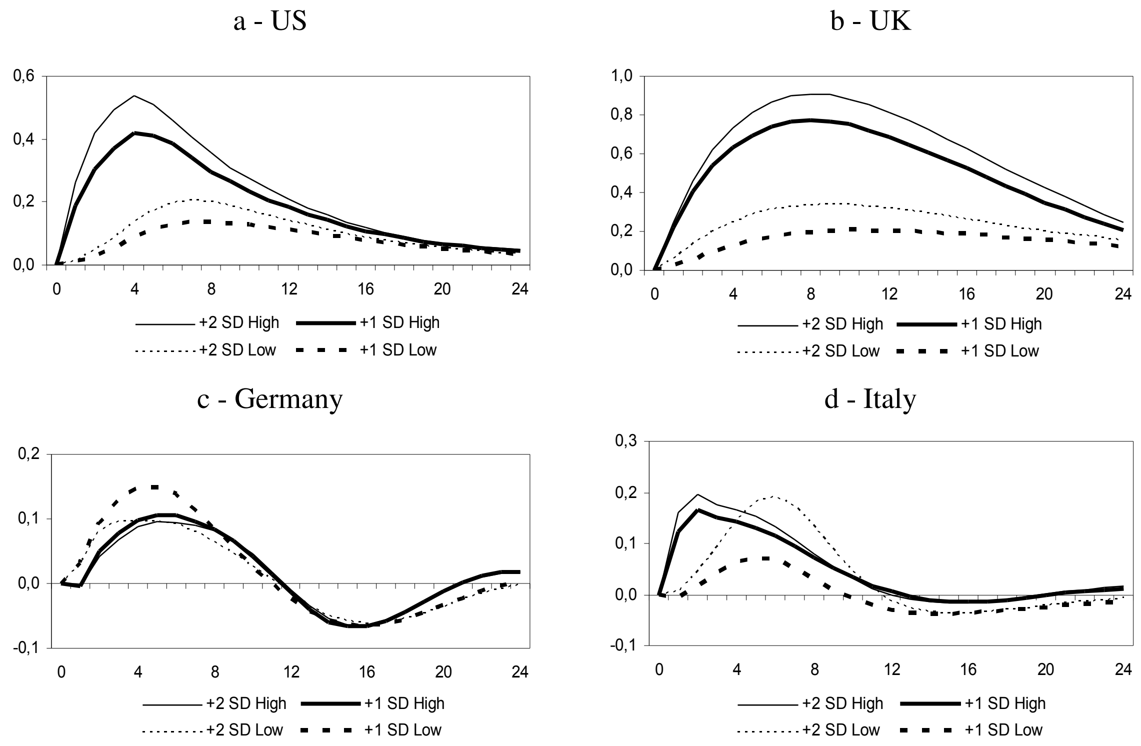
The differences among the effects of different shocks are more pronounced in the low stress regime, where the effect of a positive 2SD shock in financial stress is more than proportional in all countries. This suggests a possibility that the economy is more likely to fall into a recession after a financial stress shock, if initially a low stress regime is in place. On the contrary, output increases proportionally in response to reductions in the stress indicator in all countries except in the U.K., where the effect on output growth of an additional decrease in the financial stress measure is minor.

Table 4 – The Effects on Output Growth and on the Debt Ratio of a Shock in Financial Stress

	Financial stress regime	4 Quarters				8 Quarters			
		+2 SD	+1 SD	-1 SD	-2 SD	+2 SD	+1 SD	-1 SD	-2 SD
Output Growth									
U.S.	High	-0.366	-0.326	0.329	0.310	-0.102	-0.118	0.127	0.139
	Low	-0.208	-0.175	0.166	0.161	-0.122	-0.106	0.105	0.106
U.K.	High	-0.249	-0.193	0.191	0.137	-0.283	-0.227	0.221	0.156
	Low	-0.161	-0.100	0.081	0.063	-0.134	-0.085	0.076	0.069
Germany	High	-0.239	-0.231	0.248	0.237	-0.094	-0.091	0.135	0.132
	Low	-0.130	-0.102	0.090	0.086	-0.123	-0.116	0.106	0.104
Italy	High	-0.100	-0.104	0.158	0.140	0.040	0.030	-0.023	-0.018
	Low	-0.136	-0.085	0.092	0.082	-0.030	-0.004	0.008	0.007
Change in Debt Ratio									
U.S.	High	0.536	0.420	-0.401	-0.304	0.357	0.296	-0.290	-0.278
	Low	0.135	0.087	-0.067	-0.058	0.200	0.136	-0.124	-0.118
U.K.	High	0.735	0.632	-0.554	-0.397	0.907	0.773	-0.679	-0.531
	Low	0.249	0.123	-0.086	-0.057	0.337	0.195	-0.155	-0.133
Germany	High	0.087	0.098	-0.137	-0.140	0.081	0.083	-0.112	-0.108
	Low	0.097	0.146	-0.164	-0.175	0.064	0.086	-0.085	-0.081
Italy	High	0.166	0.142	-0.208	-0.158	0.080	0.073	-0.110	-0.109
	Low	0.146	0.061	-0.072	-0.057	0.134	0.029	-0.036	-0.028

Note: The impulse responses were rescaled to the size of the shock of one unit of FSI.

Figure 6 – Responses of Debt to a Positive Shock in Financial Stress
High and Low Regimes



Note: The impulse responses were rescaled to the size of the shock of one unit of FSI.

In line with this asymmetry in the response of output growth, the debt ratio rises after a positive shock in the FSI in both regimes (Figure 6). Generally, these increases are non-proportionally larger for a 2SD positive shock in comparison to a 1SD shock. On the other hand, improvements of the conditions in the financial markets decrease government debt, but the difference between a -1SD and a -2SD is rather small. The only exception is Germany, where the increase of the financial stress indicator by a 1SD and a 2SD causes a similar increase in debt when the economy is initially in the low stress regime. In the case of a high stress regime, the effects of financial stress shocks seem to be proportional both in sign and in size.

4.5.4 Responses over time

Nonlinear impulse responses depend not only on the estimated model coefficients but also on initial conditions, i.e. whether the economy is in the high financial stress regime at the time of the fiscal shock or not. Likewise, the impulse responses depend also on the entire history of the variables. For example, the persistence of financial stress, as well as its size, might affect the ability of fiscal policy to accomplish a switch from a high stress regime back to a low stress regime.

In the previous sections we presented the overall nonlinear impulse responses derived as the full sample average median impulse responses over both regimes. However, the fact, that nonlinear impulse responses are simulated for each point in time, allows us to investigate the time variation in the fiscal shock effects even in a model with constant parameters in the two regimes. For instance, for the U.S., the financial stress periods of the 80’s and early 90’s are associated with lower impulse responses, contrary to periods without stress (1981, 1987-1988, 1990, see Appendix 3).

Figure 7 shows the impulse responses of output growth to an initial 1 percentage point of GDP debt increase for three periods: 1981Q3-1989Q4, 1990Q1-1999Q4 and 2000Q1-2009Q4 in all countries.[11] Broadly, the effects of fiscal policy on output growth in the high financial stress regime are larger within the first two and half years, after the shock, than in the low stress regime after 2000 in all countries but the U.K. However, the initially larger effect is offset either by lower persistence of the effect (in the U.S.) or the impact on output growth becomes negative in the long term as in Germany and to some extent in Italy as well. Otherwise, the evidence of a larger positive impact of fiscal policy on output growth in times of higher stress is weak and country specific. This can also be seen from Table 7 that shows the peak multipliers corresponding to impulse responses (see Tables in Appendix 3 and Figure A3.2 illustrating the time variation with 3D plots).

In the U.S., the difference in the impact of fiscal policy on output in the low financial stress regime is negligible across periods with peak multipliers ranging from 0.182 to 0.187. With the financial stress above the threshold, the multipliers were slightly lower, below 0.180, in the 1980's and in the 1990's. Hence, the effect of fiscal policy on output growth was actually lower in periods with high stress than with low stress. In the last decade the situation was reversed and the peak multiplier reaches 0.237. However, the impact of a fiscal policy shock is not that persistent and returns to zero slightly faster in the high financial stress regime. A more detailed analysis of the simulated impulse responses in the post 2000 decade shows that fiscal policy became more effective in the periods of higher stress, which matches the 2001 recession and the 2008-2009 crisis. The peaks of the impulse responses starting in early 2001 were the largest of the entire sample.

11 The one percentage point shock was derived from the +1SD shock despite several scale effects of fiscal policy reported in previous section.

Figure 7 – Time variation of nonlinear impulse responses (high and low financial stress regimes): response of GDP to an initial 1 percentage point of GDP debt increase

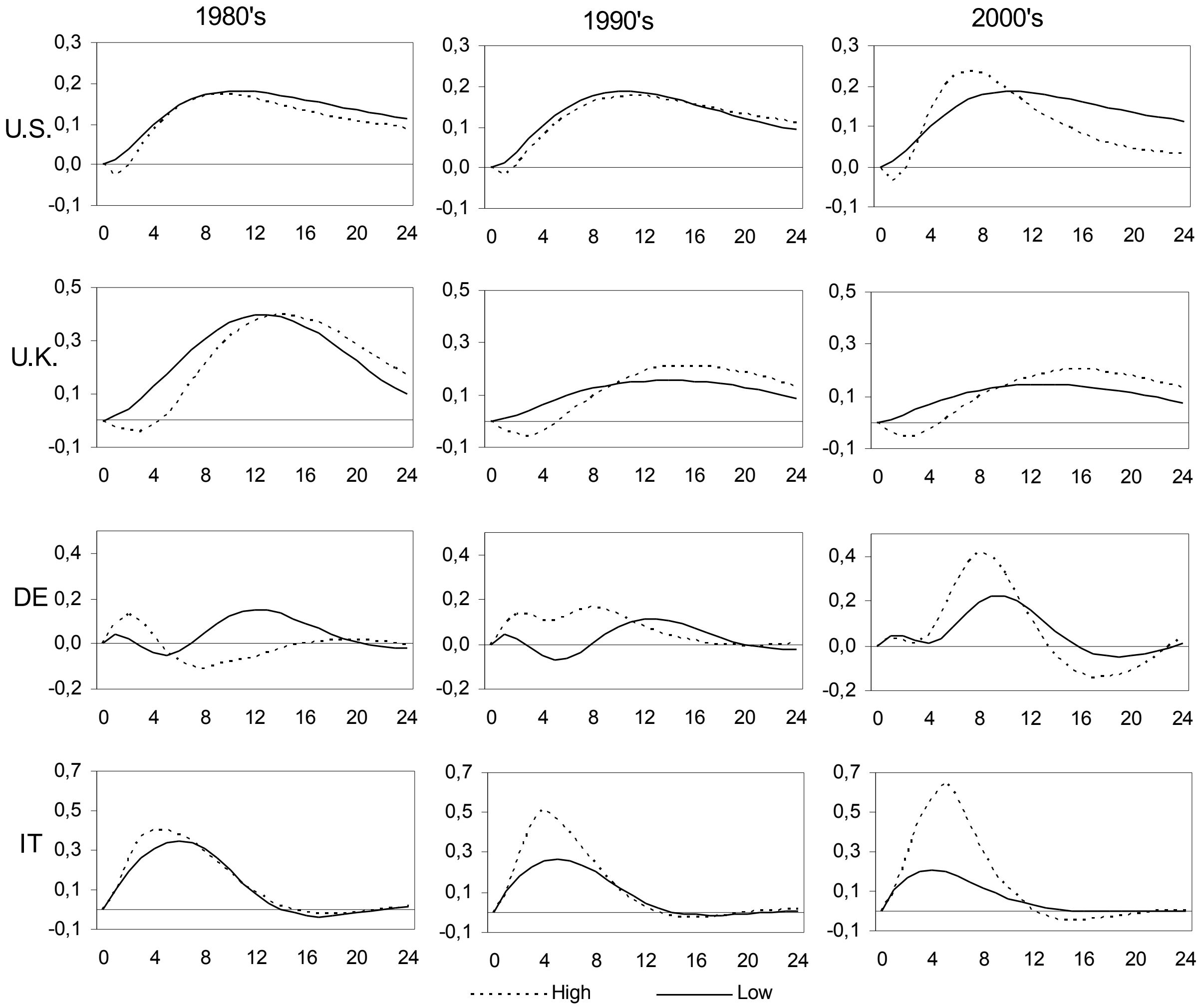


Table 5 – Peak multipliers for the response of output to a fiscal shock

	Financial stress regime	1980's	1990's	2000's
U.S.	High	0.174	0.176	0.237
	Low	0.182	0.187	0.187
U.K.	High	0.395	0.211	0.204
	Low	0.398	0.155	0.148
Germany	High	0.129	0.166	0.417
	Low	0.152	0.114	0.223
Italy	High	0.403	0.517	0.657
	Low	0.346	0.269	0.207

The sharpest fall in the size of the fiscal multiplier occurred in the U.K., where it fell in the low stress regime, from 0.398 in the eighties to 0.148 in the last decade, and from 0.395 to 0.204 in the high stress regime. We should note that this decrease in the size of the multipliers started in 1989 during a period of fiscal consolidation. In the case of the U.S. the multipliers remained lower than average during the 2008-2009 crisis as well.

In Germany, the impulse responses of output growth to a fiscal policy shock are relatively consistent in the 1980s and in the 1990s in the low stress regime. The impact was uncertain for the first two years after the shock with oscillations between positive and negative values, with a peak occurring more than three years after the shock. The high stress regime shows different patterns when comparing these two decades. In the 1980s, a fiscal policy shock increased output growth by a small margin (peaking just after two quarters at 0.129), but these positive effects quickly turned negative. On the other hand, in the 1990s the effects of a debt increase were larger and persistent. After 2000, the effect of a fiscal shock is uncertain for the first year after the shock but then it jumps up to 0.223 in the low financial stress periods and to 0.417 in the high stress periods. Such positive effect lasts up to 12-15 quarters after the shock and then both impulse responses turn into negative values.

The impact of fiscal policy on output growth has a hump-shaped pattern in Italy and it is consistent over time and across regimes. In this case, throughout time the size of the peak multiplier decreased in the low financial stress regime, from 0.346 to 0.207, and the opposite holds for the high stress regime, where it increased from 0.403 to 0.657. Regarding the 2008-2009 crisis, both Italy and Germany depict impulse responses suggesting larger than average impacts on output growth in 2008 but in 2009 sizes of fiscal multipliers decreased.

4.6 Conclusion

We analyzed the interactions between fiscal and financial developments in times of financial instability. The effects of fiscal policy were estimated using a threshold VAR with macro, fiscal and financial variables and with regime switching determined by a measure of financial stress. The application of a nonlinear framework with regime switching was motivated by the debate on the ability of fiscal policy to shorten recessions and to facilitate a subsequent recovery, and its empirical adequacy was confirmed by formal nonlinearity test in TVAR model. Furthermore, the identified periods of financial stress are characterised by lower growth and in a number of cases coincide with recessions.

Unlike their linear counterparts, nonlinear impulse responses are differences between the simulated paths of endogenous variables with and without an initial shock, either in fiscal policy or in financial stress conditions. Given its nature, this approach allows to take into account future regime switches caused by a shock on any endogenous variable and not only on financial stress, which determines the alternative regimes in our model. The other advantage is that the framework of nonlinear impulse responses can be used to recover time variance in impulse responses.

The empirical results and the implications of our model are threefold. First, the differences among the fiscal multipliers of various sizes and signs of shocks are small in all countries. However, the initial state of the economy matters and both multipliers and the estimated responses to fiscal shocks differ across regimes. The difference between the high and low financial stress regimes is lowest in the U.S. Moreover, fiscal policy shocks have a larger effect on output growth in both Euro area countries, Germany and Italy, although the multipliers for Germany are lower. On the contrary, for the U.K. the multipliers are much lower in the high financial stress regime.

Specifically, for the U.S. cumulative multipliers are between 0.45-0.46 with minor differences between signs and sizes of shock in both regimes. On the other hand, the response of output to fiscal shocks is faster in high stress regime in comparison to the low stress regime and the peak is reached earlier by two quarters. The U.K. has the lowest effects of a fiscal shock on output growth in the high stress regime, with the cumulative multiplier over three years being between 0.22 and 0.3. Interestingly, if the fiscal shocks occur in the low financial stress regime, the cumulative multipliers are around 0.50-0.54. The multipliers are largest in Italy with a size of the cumulative multiplier after three years of about 0.82-0.87 for the high stress regime, and 0.48-0.49 for the low stress regime. In Germany the cumulative fiscal multiplier is 0.3 in the high stress regime and almost zero when the economy is initially in the low stress regime, implying strong crowding-out effects in the economy.

Second, the ability of fiscal policy to affect output growth evolved over time. Indeed, the fiscal multipliers increased since the 1990's in the high financial stress regime in all countries except the U.K. where they remained stable. The multipliers associated to the responses with initial conditions in the low financial stress regime decreased over time in the U.K. and in Italy, remained stable in the U.S., and increased in Germany.

Third, financial stress shocks have strong negative effects on output growth and such effects are also nonlinear. The negative effect is largest in the high stress regime, but it is still rather proportional, and the difference between small and large increases of financial stress is small. In the low stress regime, output growth falls much more in response to a large increase in financial stress suggesting an increased probability of a shift in the regime.

Therefore, we have found evidence of nonlinearities in the effects of a fiscal shock depending on the initial conditions, determined by the existence of financial stress, diverse levels of government indebtedness, and, of course implicitly assumed different monetary policy behaviour. In addition, both the multipliers and the nature of these nonlinearities vary across countries and evolve over time. Finally, the estimated thresholds also match economic recessions, and the effectiveness of fiscal policy in the context of different financial stress regimes also differs across country, naturally something to bear in mind by policy makers.

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Appendix

Appendix 1. Data description and sources

Variables in Threshold VAR

y_t GDP, annual growth rate of the log of the real GDP (*Y*) used: $y_t = \log(Y_t) - \log(Y_{t-4})$.

p_t Price level (P), annual growth rate of logs used: $p_t = \log(P_t) - \log(P_{t-4})$.

i_t Short-term interest rate.

f_t Annual change in the debt to GDP ratio: $f_t = D_t - D_{t-4}$.

s_t Financial stress index.

Financial stress variables

FSI (sum of subsequent components).

Bank stress (normalized beta of stocks of banking sector + normalized TED spread + normalized inverted term structure).

Stock market stress (volatility of stocks + returns of stock + spread of corporate bonds, all normalized).

Exchange rate volatility.

Data Sources

United States

Nominal GDP: IMF IFS (IFS.Q.111.9.9B.B$C.Z.F.$$$).

GDP deflator: IMF IFS (IFS.Q.111.9.9B.BIR.Z.F.$$$).

Interest rate: Federal funds rate, FRED, series FEDFUNDS.

Government debt: Federal Debt held by the Public, FRED, series FYGFDPUN.

Stock prices: Dow Jones Industrial Index, quarterly averages.

TED spread: Spread between treasury bills rate (3M) and interbank interest rate represented by the Eurodollar 3M rate, IMF IFS.

United Kingdom

Nominal GDP: IMF IFS (IFS.Q.112.9.9B.B$C.Z.F.$$$), rolling sum of 4 quarters to calculate the annual GDP.

GDP deflator: IMF IFS (IFS.Q.112.9.9B.BIR.Z.F.$$$).

Interest rate: End of quarter Sterling interbank lending rate, 1 month, average; Bank of England, series IUQVNEA.

Government debt: Since 2000 Quarterly Government Debt (Maastricht Debt) for General Government, Eurostat; older data from other sources, merged using growth values in overlapping periods (Public sector debt, National Statistics, series BKQK; Quarterly amounts outstanding of General Government sterling and all foreign currency consolidated gross debt total (in sterling millions), Bank of England, series DPQG004).

Stock prices: Financial Times Stock Exchange (FTSE) 100 Index - Historical close, end of period, UK pound sterling, provided by DataStream.

TED spread: Spread between treasury bills rate (3M) and interbank interest rate represented by LIBOR 3M rate, IMF IFS.

Germany

Nominal GDP: Federal Statistical Office, DeStatis, National Accounts, Gross Domestic Product since 1970, Quarterly and Annual Data. The time series before the German Unification was rescaled to the post-unification period using growth rates of quarterly data that overlap in 1991. The GDP deflator was calculated as the ratio of nominal and real GDP (available as index of 2000=100 only), rescalled to the post unification period using quarterly growth rates as well.

CPI: IMF IFS (IFS.Q.134.6.64).

Interest rate: Money market rates reported by Frankfurt banks, monthly average of overnight money.

Government debt: Statistische Angaben: Umrechnungsart: Endstand, Euro, Millionen, Bundesbank. SeriesBQ1710, BQ1720, Central, state and local government debt; Total debt (excluding hospitals).

Stock prices: Share prices, IMF IFS.

Italy

Nominal GDP: OECD (OEO.Q.ITA.GDP).

GDP deflator: IMF IFS (OEO.Q.ITA.PGDP).

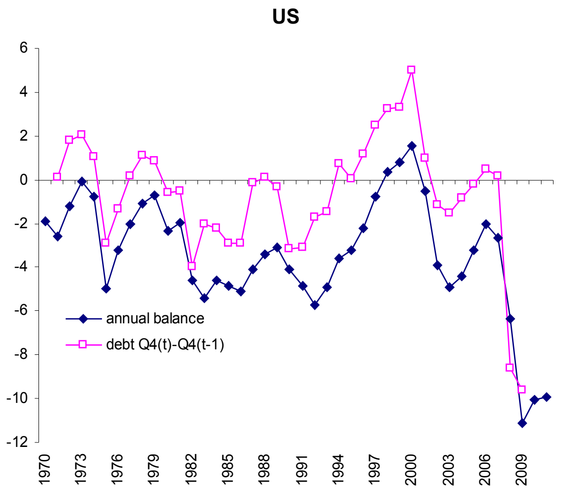
Interest rate: money market rate, IMF IFS.

Government debt: General Government debt, Banca d'Italia.

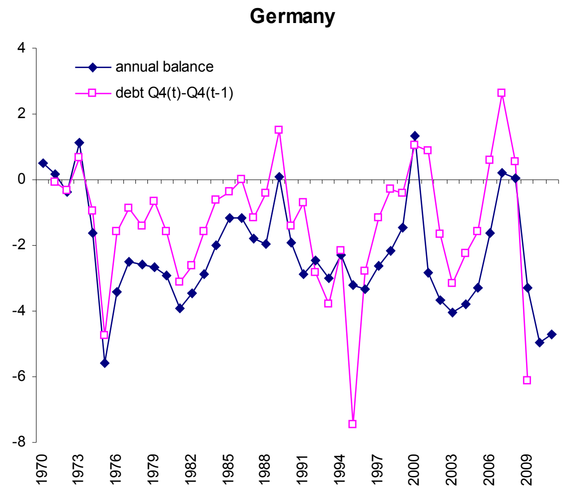
Stock prices: Share prices, IMF IFS.

Figure A1 – Government debt and budget balance ratios

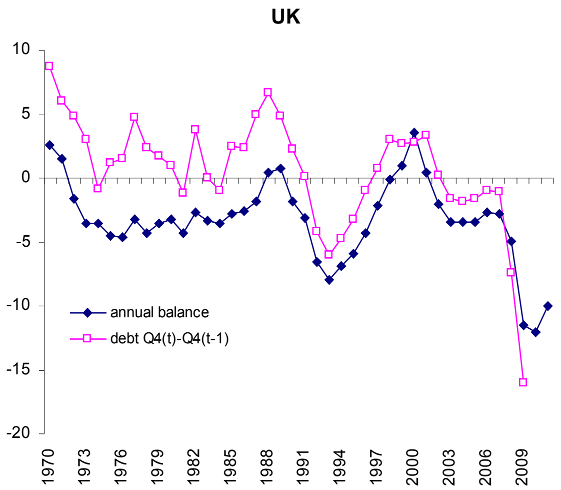
A2.1



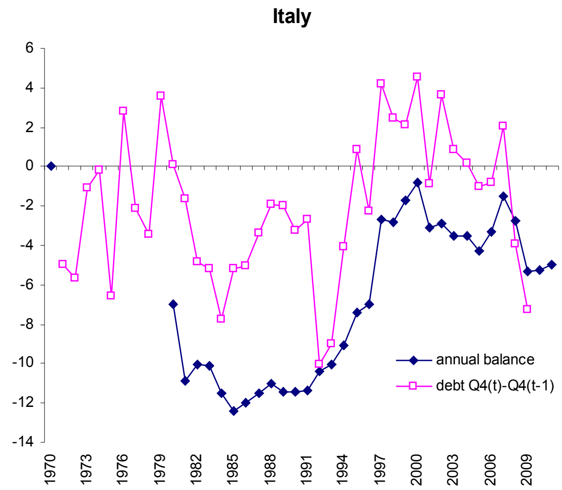
A2.2



A2.3



A2.4



Note: inverted scale for the change in the debt ratio.
Source: AMECO database (annual budget balance data) and national central banks (quarterly government debt data).

Appendix 2: Nonlinear impulse response functions

The algorithm for computing nonlinear impulse response functions (NIRF) follows Koop, Pesaran and Potter (1996). NIRF is defined as the difference between the forecasts of variables with initial shock ε_t and without initial shock into a variable of interest.

Formally, the nonlinear impulse response functions are defined as

$$NIRF_y = (k, \varepsilon_t, \Omega_{t-1}) = E(Y_{t+k} | \varepsilon_t, \Omega_{t-1}) - E(Y_{t+k} | \varepsilon_t^{0,} \Omega_{t-1}), \qquad \text{(A.1)}$$

where Y_{t+k} is a vector of variables at horizon k, Ω_{t-1} is the information set available before the time of shock ε_t. NIRF are computed by simulating the model with and without the shock. Note that the ε_t^0 denotes stochastic disturbance at time 0 that would occur under the no-shock scenario. The algorithm proceeds as follows:

1) Pick a history Ω_{t-1}.
2) The shocks for the periods from 0 to q are drawn from the residuals of the estimated VAR model.
3) For each initial value this sequence of shocks is fed through the model to produce forecasts conditional on initial conditions.
4) Repeat step 2) with the initial shock into one variable equal to +/- 1 or 2 SD to get forecasts if there was an initial shock.
5) The difference between the forecasts from step 2 and 3 is the impulse response function. Repeat this 500-times and derive an average impulse response for this particular initial condition
6) Repeat steps 2-4 for each initial conditions. Final impulse responses are average impulse responses over initial conditions of each regime. Confidence bands derived from quantiles of empirical distribution. We use a 50% confidence bands here.

Appendix 3: Additional tables and figures

Table A3.1 – Multipliers at selected horizons, for the response of output to a fiscal shock

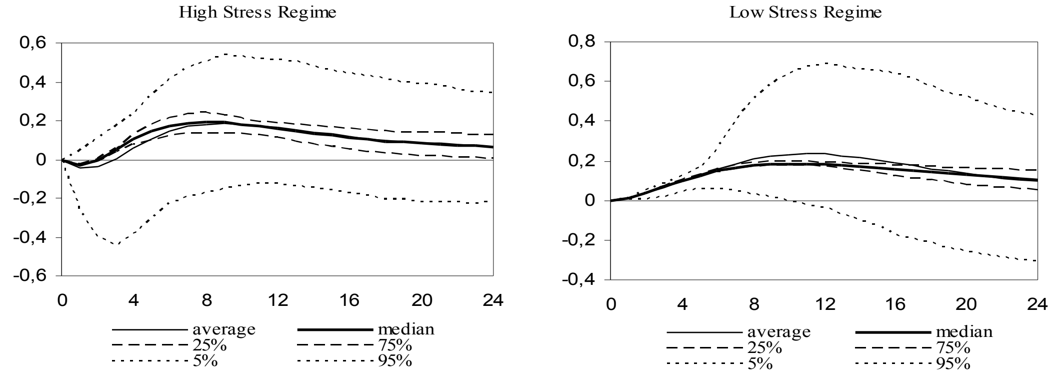
	Regimes	1981Q3:1989:Q4			1990Q1:1999Q4			2000Q1:2009Q4		
		4Q	8Q	12Q	4Q	8Q	12Q	4Q	8Q	12Q
U.S.	High	0.084	0.168	0.160	0.076	0.161	0.175	0.139	0.233	0.145
	Low	0.099	0.172	0.180	0.101	0.177	0.184	0.100	0.178	0.184
U.K.	High	-0.020	0.211	0.375	0.076	0.161	0.175	-0.030	0.096	0.176
	Low	0.126	0.307	0.398	0.101	0.177	0.184	0.066	0.124	0.145
Germany	High	0.034	-0.110	-0.058	0.108	0.166	0.075	0.044	0.417	0.116
	Low	-0.041	0.046	0.152	-0.049	0.003	0.112	0.013	0.195	0.162
Italy	High	0.403	0.294	0.083	0.517	0.240	0.024	0.573	0.294	0.003
	Low	0.307	0.310	0.076	0.257	0.204	0.045	0.207	0.116	0.029

Table A3.2 – Peak multipliers, for the response of output to a fiscal shock: the 2008-2009 period

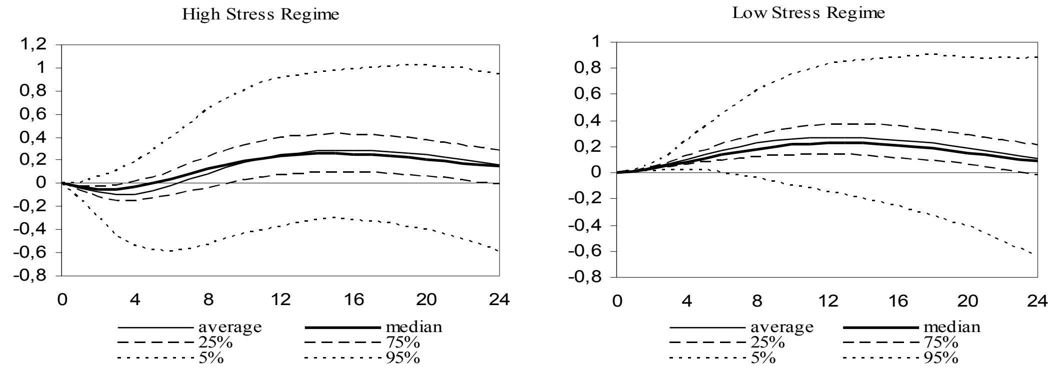
	2008Q2	2008Q3	2008Q4	2009Q1	2009Q2	2009Q3	2009Q4	average
U.S.	0.268	0.289	0.263	0.230	0.200	0.181	0.180	0.189
U.K.	0.286	0.221	0.212	0.217	0.187	0.121		0.279
Germany	0.392	0.367	0.342	0.228	0.192	0.185	0.212	0.139
Italy	0.810	1.031	1.097	0.536	0.308	0.228		0.505

Figure A3.1 – Distribution of the Response of Output to Fiscal Policy Shock, 1% of GDP)

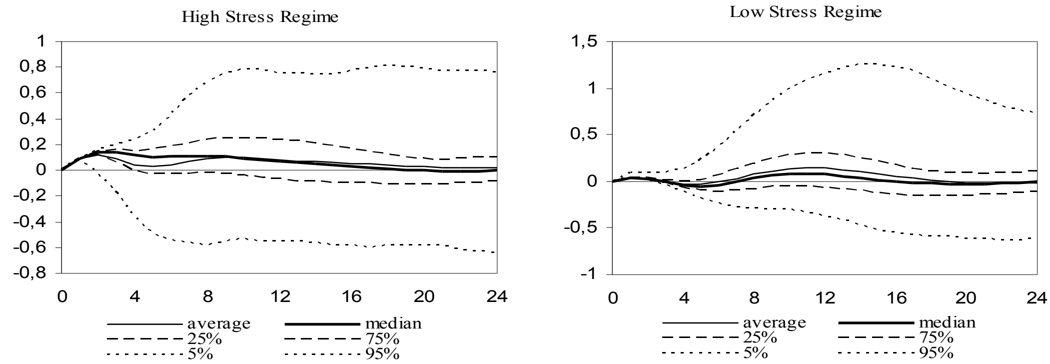
U.S.



U.K.



Germany



Italy

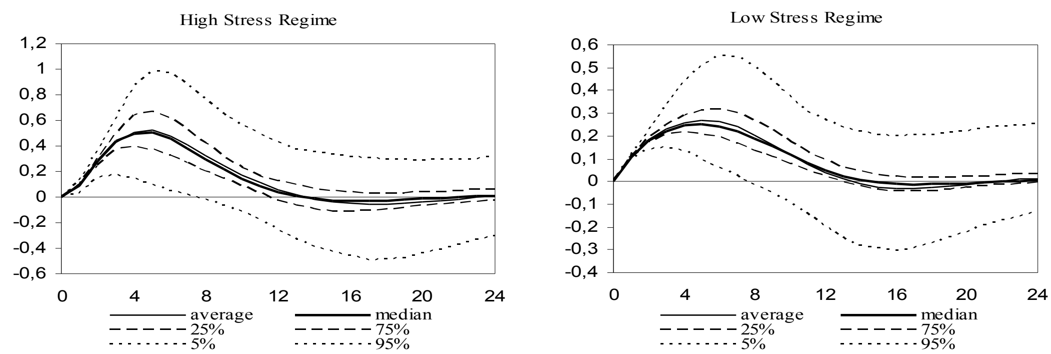


Figure A3.2 – Median impulse responses of 1% fiscal shock over time