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

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We examine whether and how selected central banks responded to episodes of financial stress over the last three decades. We employ a new monetary-policy rule estimation methodology which allows for time-varying response coefficients and corrects for endogeneity. This flexible framework applied to the USA, the UK, Australia, Canada, and Sweden, together with a new financial stress dataset developed by the International Monetary Fund, not only allows testing of whether central banks responded to financial stress, 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. Our 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.

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Nontechnical Summary

The recent financial crisis has intensified the interest in exploring in greater detail the nexus between monetary policy and financial stability. Although keeping the financial system stable is a major task, often delegated to central banks, how to consider financial stability concerns for monetary policy decision-making remains a puzzling question. Monetary policy is likely to react to financial instability in a non-linear way. 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.

This paper examines the reactions of the main central banks (the US Fed, the Bank of England, the Reserve Bank of Australia, the Bank of Canada, and Sveriges Riksbank) during periods of financial stress over the last three decades. In particular, we estimate the time-varying policy rule of each bank to assess whether and how its policy rate was adjusted in the face of financial instability. We track financial stress by means of a continuous financial stress indicator developed recently by the International Monetary Fund as well as its main subcomponents (banking stress, stock-market stress, and exchange-rate stress). Therefore, our empirical framework is suitable for detecting the periods and types of stress that were perceived as the most worrying and for quantifying the intensity of the policy response.

Although theoretical studies disagree about the viability of considering financial instability for interest-rate setting, our empirical results suggest that central banks often alter the course of monetary policy in the face of high financial stress, mainly by decreasing their policy rates. Yet the size of this response varies substantially over time and across countries as well as in terms of specific types of financial stress. The recent financial crisis evoked the most generalized response and interest rates decreased in the range of 50 to 200 basis points owing solely to financial instability concerns (above what could be attributed to the decline in inflation expectations and output below its potential). However, significant interest rate adjustment to financial instability has been found for all central banks over the whole sample period. In terms of the specific components of stress, most banks seemed to respond to stock-market stress and bank stress, while exchange-rate stress drove central bank reactions only in more open economies.

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.

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 central banks reacted to periods of financial instability and, in particular, whether and how the interest-rate-setting process 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 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

[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.

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.

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.

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

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

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 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 an NK DSGE model with credit friction to evaluate the performance of alternative policy rules that are augmented by a response to credit spreads and 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.

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)

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

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.

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. Data and Empirical Methodology

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

[5] The IMF Financial Stress Index has recently been applied by Melvin and Taylor (2009) to analyze exchange rate crises.

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.[7] On the one hand, 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.[8]

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.[9] 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.[10] 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).

[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.

[7] The experience from the recent crisis represents an evident drawback of using interbank rates as a proxy for monetary policy intentions. In particular, the interbank markets froze at longer maturities (3–12 months) and the term premium increased sharply in most countries. In addition, in open economies, central banks lost control over very short maturities. For example, EONIA has been below the ECB repo rate for an extended period as major financial institutions can borrow more cheaply in the yen or dollar markets.

[8] 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.

[9] 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.

[10] 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.

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);

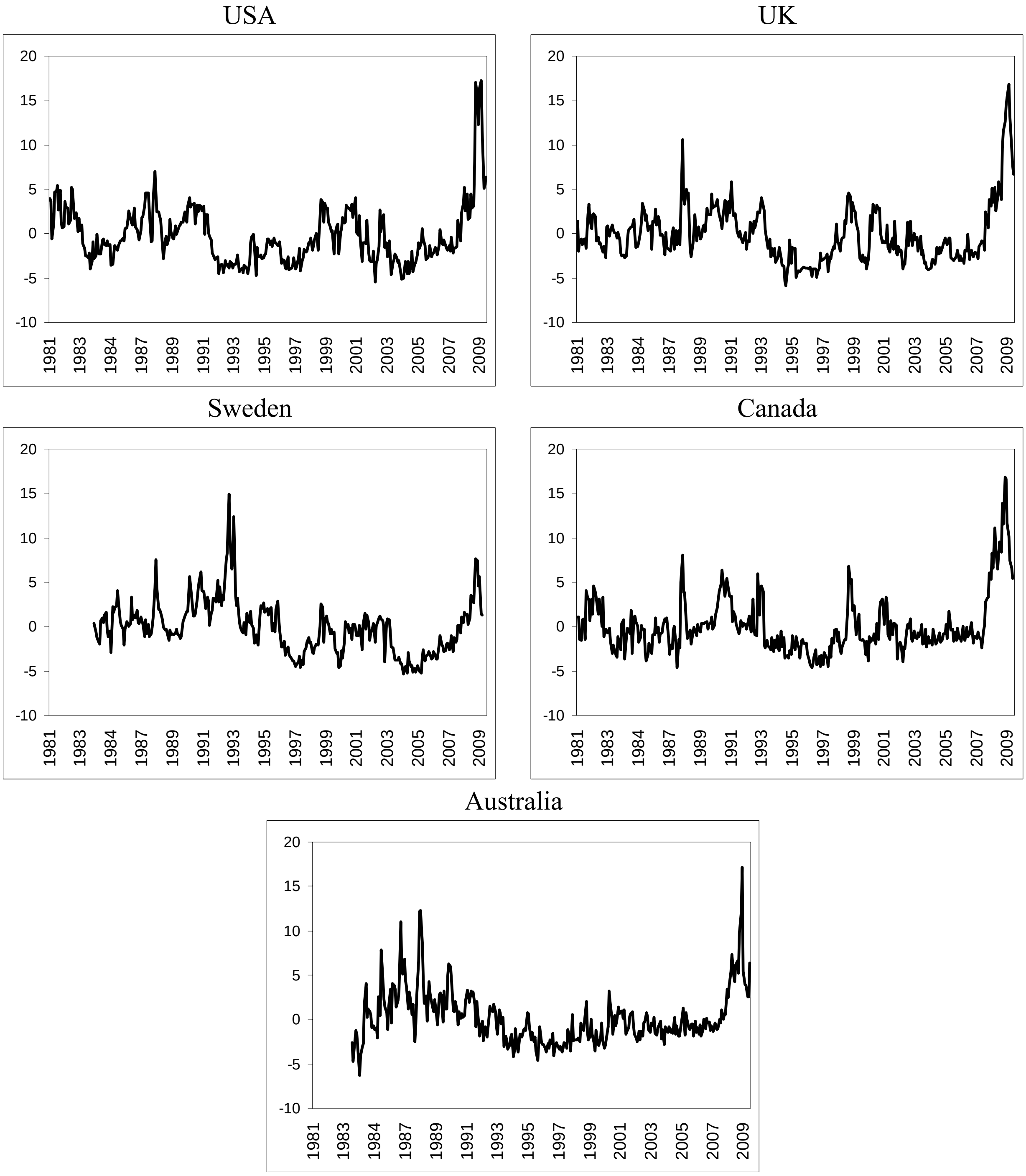
(ii) Securities-market-related sub-index components: corporate bond spread (the difference between corporate bonds and long-term government bond yields), stock-market returns (monthly returns multiplied by -1), time-varying stock-return volatility from the GARCH(1,1) model;

(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.

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.

Figure 1: IMF Financial Stress Indicator



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

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:

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

where r_t^* denotes the targeted interest rate, $\bar{r}$ is the policy neutral rate[11], 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 π_{t+i}^* is the central bank's inflation target.[12] y_{t+j} 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:

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

where $\rho \in \left[0,1\right]$ 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), assuming rational expectations we replace unobserved forecast variables by their actual realization (this includes endogeneity in the model, which is dealt with by instrumental variable estimation; see the details below), pass the inflation forecast to the generic intercept α, and include the measures of financial stress described above, which results in Eq. (3):

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

While in Eq. (1) the term α coincides with the policy-neutral rate $\bar{r}$, its interpretation is not straightforward once the model is augmented by additional variables. Note that the financial

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

[12] 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.

stress index x_{t+k} does not appear within the square brackets because it is not a variable that determines the target interest rate r_t^*, 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, *j* equal to 0 and *k* equal to -1.[13] 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. Indeed, regime change has been a significant feature of monetary policy conduct over recent decades. 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.[14] 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.

[13] 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*. 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 of statistical methods. 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 be preemptive rather than reactive.

[14] 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).

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

$$r_t = (1-\rho_t)\left[\alpha_t + \beta_t(\pi_{t+i}) + \gamma_t y_{t+j}\right] + \rho_t r_{t-1} + \delta_t x_{t+k} + \varepsilon_t \tag{4}$$

$$\alpha_t = \alpha_{t-1} + \vartheta_{1,t},\ \vartheta_{1,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_1}\right) \tag{5}$$

$$\beta_t = \beta_{t-1} + \vartheta_{2,t},\ \vartheta_{2,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_2}\right) \tag{6}$$

$$\gamma_t = \gamma_{t-1} + \vartheta_{3,t},\ \vartheta_{3,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_3}\right) \tag{7}$$

$$\delta_t = \delta_{t-1} + \vartheta_{4,t},\ \vartheta_{4,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_4}\right) \tag{8}$$

$$\rho_t = \rho_{t-1} + \vartheta_{5,t},\ \vartheta_{5,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_5}\right) \tag{9}$$

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

$$y_{t+j} = Z'_{t-m}\psi + \sigma_v v_t,\ v_t \sim i.i.d.N(0,1) \tag{11}$$

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

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.[15] 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:

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

π_{t-1}, π_{t-12} (π_{t-4} for CAN and SWE), y_{t-1}, y_{t-2}, r_{t-1}, 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, v_t, and ι_t and the error term ε_t is $\kappa_{\varphi,\varepsilon}$, $\kappa_{v,\varepsilon}$, and $\kappa_{\iota,\varepsilon}$, respectively (note that σ_φ, σ_v, and σ_ι are the standard errors of φ_t, v_t, and ι_t, respectively). Consistent 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, v_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 ζ_t is uncorrelated with the regressors.

$$r_t = (1-\rho_t)\left[\alpha_t + \beta_t \pi_{t+6} + \gamma_t y_{t-1}\right] + \rho_t r_{t-1} + \delta_t x_{t-1} + \kappa_{\varphi,\varepsilon}\sigma_\varepsilon \varphi_t + \kappa_{v,\varepsilon}\sigma_\varepsilon v_t + \kappa_{\iota,\varepsilon}\sigma_\varepsilon \iota_t + \zeta_t$$
$$\zeta_t \sim N\left(0, (1-\kappa^2_{\varphi,\varepsilon} - \kappa^2_{v,\varepsilon} - \kappa^2_{\iota,\varepsilon})\sigma^2_{\varepsilon,t}\right) \tag{13}$$

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 residuals $\sum_{t=1}^{T}\zeta^2$, uses minimization of the weighted sum of the squares:

$$\sum_{t=1}^{T}\zeta^2 + \theta_1 \sum_{t=1}^{T}\vartheta_1^2 + \theta_2 \sum_{t=1}^{T}\vartheta_2^2 + \ldots + \theta_{\mathrm{n}} \sum_{t=1}^{T}\vartheta_{\mathrm{n}}^2 \tag{14}$$

where the weights θ_i are the inverse variance ratios of the regression residuals ε_t and the shocks in time-varying coefficients ϑ_t, that is, $\theta_i = \sigma^2/\sigma_i^2$. This approach balances the fit of the model and parameter stability. Additionally, the time averages of the regression coefficients, estimated by a weighted least squares estimator, are identical to their 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;

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

- 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 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 ϑ_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

[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 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.

deviation from the target interest rate, as implied by the macroeconomic variables, due to the response to financial stress.

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. 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.

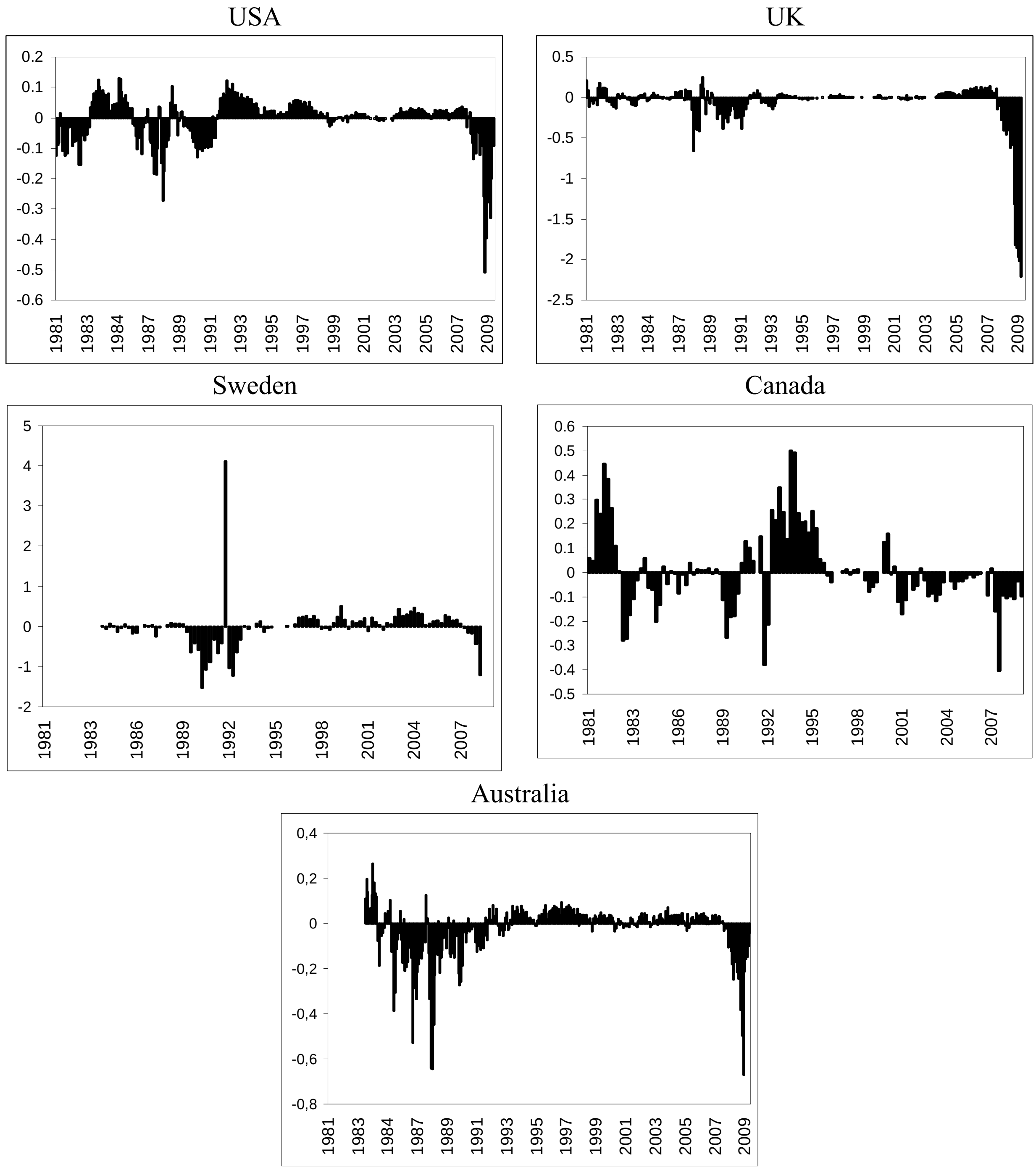
4.1 Financial-Stress Effect

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.

[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.

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 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 preemptive reaction 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.

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.

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[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.

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.