Three Essays in International Macroeconomics

Inaugural-Dissertation

zur Erlangung des Grades eines Doktors
der Wirtschafts- und Gesellschaftswissenschaften

durch

die Rechts- und Staatswissenschaftliche Fakultät der
Rheinischen Friedrich-Wilhelms-Universität Bonn

vorgelegt von

THOMAS HETTIG

aus Hannover

Bonn, 2019

Dekan:	Prof. Dr. Jürgen von Hagen
Erstreferent:	Prof. Dr. Gernot Müller (Universität Tübingen)
Zweitreferent:	Prof. Dr. Benjamin Born (Frankfurt School of Finance & Management)
Tag der mündlichen Prüfung:	27.03.2019

Acknowledgments

This thesis would not have been possible without the support of many people. I first and foremost would like to thank my main supervisor, Gernot Müller, for his guidance, encouragement and faith in my abilities. I learned a great deal from him, especially in the areas of research and economic thinking. He also enriched my leadership, writing and presentation skills. It was a great honor for me when Petr Sedlacek assured me (after presenting in Tübingen) that I was indeed Gernot's student since I asked the same questions as he did.

I am also very grateful to my second supervisor Benjamin Born for his availability and guidance, especially in the econometric parts of this thesis. His input had a large impact on this thesis and improved the chapters significantly. Further, I wish to thank Keith Kuester for shaping my thinking during the doctoral studies. It was very helpful that he always challenged us to understand the main economic mechanisms of research papers. I also learned much from his questions in the weekly research seminars.

Furthermore, I want to thank the administration of the Bonn Graduate School of Economics (BGSE) for creating such a nurturing environment that allowed me to write this thesis. In this regard I would like to highlight the assistance and efforts of Silke Kinzig and Britta Altenburg. I also had the privilege of becoming a research fellow at the Max Planck Institute (MPI) for Research on Collective Goods in Bonn. I benefited from discussions with Martin Hellwig during my time at the MPI. I also wish to thank Heidi Morgenstern and Monika Stimpson for their administrative support. I finished this thesis at the University of Tübingen and benefited from discussing my research with Willi Kohler, Christian Manger, Thomas Dimpfl, and Jantje Sönksen. I am indebted to Martina Mayer for her administrative support. Overall, I am grateful for the financial support of the BGSE, the MPI, the German Science Foundation (DFG) and the University of Tübingen.

My doctoral colleagues also had a significant impact on this thesis. I very much appreciated the advice of and discussions with Yao Chen, Martin Wolf, Alex Scheer, Dmitry Kuvshinov, Patrick Hürtgen, Peter Eppinger, Francesco D'Ascanio, and Susanne Wellmann. In this context I also wish to thank Marc Faupel for the fruitful discussions. My thanks extend to my colleagues at the BGSE, the MPI and at University of Tübingen for the great time I had with them.

Part of this thesis was conducted while I visited the research department of the Brazilian Central Bank in Brasília. I would like to express my gratitude to the department for their hospitality. Especially, I want to thank Sérgio Lago for his effort—which made it possible for me to visit the department—and for discussing my research with him.

This thesis would not have been possible without the exceptional support of my parents. Their help cannot be overestimated. I benefited a lot from their encouragement and from the fact that they were always available in times of need. Thanks also to my extended family for all the support and love given to me.

Finally, I want to thank my God (Father, Son, and Holy Spirit) for his continual support, love, and grace towards me. In this regard I also wish to thank my friends from church. Although it is impossible to name all of them, I will name only a few: I am very thankful to Dorothea Bastgen, Micha Fetzner, Immanuel Wahl, Ben Rössler, Andrea Rössler, and Mareile Janko for their support and encouragement.

Contents

List of Figures

List of Tables

Introduction

This thesis challenges three views in international macroeconomics with important policy implications. The first chapter asks how fiscal policy should be designed in a monetary union and draws novel conclusions. The second chapter challenges the notion how nominal exchange rates respond to exogenous interest rate changes under rational expectations. It provides new empirical evidence and a theoretical explanation of this evidence. The third chapter calls the view into question that the German current account surplus is too large by rationalizing the surplus on the basis of a theoretical model. In what follows, I provide a short summary of the main findings of each chapter.

Chapter 1 is based on joint work with Gernot Müller. In it we analyze to which extent fiscal policy should be coordinated in a monetary union. According to the pre-crises consensus there are separate domains for monetary and fiscal stabilization in a currency union. While the common monetary policy takes care of union-wide fluctuations, fiscal policies should be tailored to meet country-specific conditions. This separation is no longer optimal, however, if monetary policy is constrained by an effective lower bound on interest rates. Specifically, we show that in this case there are benefits from coordinating fiscal policies across countries. By coordinating fiscal policies, policymakers are better able to stabilize union-wide activity and inflation while avoiding detrimental movements of a country's terms of trade.

Chapter 2 is based on joint work with Gernot Müller and Martin Wolf. In it we revisit Dornbusch (1976)'s exchange rate overshooting hypothesis. First, we use local projections to estimate the response of the effective exchange rate of the US dollar to US monetary policy shocks. Following a monetary contraction, the dollar undershoots: it appreciates on impact, but less so than in the long run. Second, we develop and estimate a New Keynesian model with information frictions. Market participants do not fully observe the natural rate and attach probability to the scenario that the monetary policy shock represents in fact an endogenous response to movements in the natural rate. As they learn over time the true nature of the shock, the exchange rate continues to appreciate, as in the data.

Chapter 3 is based on joint work with Marc Faupel. In it we test whether the so called intertemporal approach to the current account is able to explain German current account data. For this purpose, we perform established present value tests of the intertemporal model. We find that the cross-equation restrictions of the model cannot be rejected and that the large current account surplus is predicted by the intertemporal model. Hence, according to this benchmark the German current account surplus can be justified by the intertemporal approach.

I hope that this thesis contributes to our understanding about how the economy works and that it provides policymakers with helpful insights for their decision making.

Chapter 1

Fiscal policy coordination in currency unions at the effective lower bound

Joint with Gernot Müller

1.1 Introduction

In the wake of the global financial crisis, fiscal policy staged a comeback as a stabilization tool. Figure 1.1 displays two rough measures of the discretionary fiscal stance, both for the US and the euro area. The left panel shows the change of the cyclically adjusted government budget deficit, measured in percentage-point changes relative to the pre-crisis year 2007. The right panel shows the level of government consumption relative to trend output. Both measures are indicative of an expansionary fiscal stance during the recession: deficits rose sharply after 2007, as did government spending. It appears, however, that fiscal stabilization has been used more timidly in the euro area: not only did deficits increase less than in the US, government spending was also raised relatively less, given its higher pre-crisis level.

One possible explanation is that euro-area fiscal policy is largely determined at the country level, rather than at the union level and, hence, there may have been a failure to coordinate fiscal stabilization across the member states of the euro area.[1] In line with this conjecture, there have been calls for stronger policy coordination, urging European governments to engineer a larger fiscal expansion during 2008–09 (see, for instance, Krugman, 2008). For the same reason, the shift to austerity in the euro area after 2010 may have been excessive, as argued by many observers (see, for instance, Cotarelli, 2012). Against this background, we ask whether fiscal-stabilization policies should be coordinated across the

[1] To be sure, there has been an attempt to coordinate European fiscal stabilization policies, namely through the *European Economic Recovery Plan*, discussed and legislated in 2008–09. According to Cwik and Wieland (2011) the measures foreseen by the plan amounted to 1.04 and 0.86 percent of 2009 and 2010 GDP, respectively. Hence, they were considerably smaller than those due to US legislation under the *American Recovery and Reinvestment Act* which amounted to roughly 5 percent of GDP.

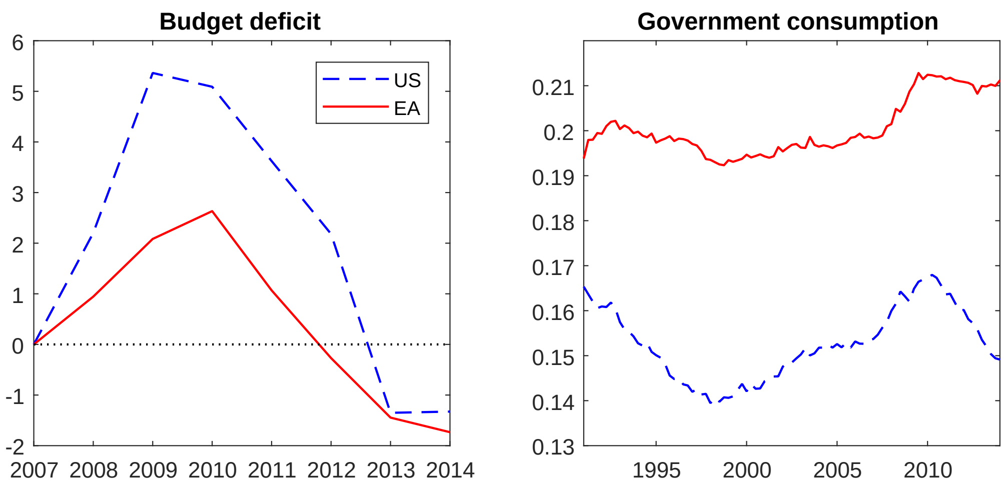

Figure 1.1: Cyclically adjusted deficit (annual observations, change relative to 2007, measured in percentage points of potential output) and government consumption of general government (in units of trend output, quarterly observations); solid lines: Euro area (EA), dashed lines: United Staates (US); source: OECD Economic Outlook.

member states of a currency union with a view towards stabilizing area-wide activity and inflation.

According to the pre-crises consensus fiscal stabilization should be geared towards country-specific conditions, because the common monetary policy can take care of union-wide fluctuations (Beetsma and Jensen, 2005; Kirsanova et al., 2007; Galí and Monacelli, 2008).[2] The recent economic and financial crises have exposed a shortcoming of this paradigm: in a severe economic downturn monetary policy may be constrained by an effective lower bound (ELB) on nominal interest rates and thus be unable to stabilize fluctuations at the union level. Moreover, it is precisely under these circumstances that fiscal policy is very effective in stabilizing economic activity (Christiano et al., 2011; Woodford, 2011).

In our analysis we therefore explicitly account for the possibility that an ELB constrains

[2]Earlier contributions also allow for the possibility that the objectives of monetary and fiscal policy differ. This does not necessarily strengthen the case for coordination (Dixit and Lambertini, 2003). In fact, fiscal coordination may even be harmful (Beetsma and Bovenberg, 1998). Dixit and Lambertini (2001) offer some qualifications as well as further references.

monetary policy. We do so within the framework of Galí and Monacelli (2008). It specifies a currency union which consists of a continuum of small open economies, each negligible in terms of aggregate outcomes. Yet as countries specialize in the production of a specific set of goods, domestic policies—if enacted unilaterally—will generally impact a country's terms of trade. In the absence of policy coordination, the optimal policy will therefore be conducted with a view towards its effect on the terms of trade. Instead, by coordinating on a common policy, countries can internalize this "terms-of-trade externality".[3] Hence, optimal policies will generally differ depending on whether there is coordination across countries or not.

In terms of country-specific policies, we focus on government spending. We assume that the government purchases only domestically produced goods, financed by taxes levied on domestic households. We consider a representative household in each country which supplies labor and trades a complete set of state-contingent assets across countries. Its consumption basket includes goods produced in all countries of the union, but is biased towards domestically produced goods. Goods, in turn, are produced in a monopolistic competitive environment and firms are restricted in their ability to adjust prices. In "normal" (or pre-crisis) times monetary policy is able to perfectly stabilize inflation and output at the union level: there is no need for fiscal coordination across countries. We contrast this situation with a "crisis scenario" where monetary policy is unable to lower interest rates sufficiently in response to a union-wide contractionary shock, because it is constrained by an ELB.[4]

We determine the optimal discretionary adjustment of government spending in the crisis scenario. Under coordination fiscal policies are set to maximize union-wide welfare. In the absence of coordination each fiscal policymaker maximizes country-specific welfare. We find that—in line with the conjecture above—countries provide too little stimulus at the ELB in the absence of coordination. Intuitively, local policymakers are keen to avoid the terms of trade appreciating too much with higher spending, as this lowers the demand for domestically produced goods at times of economic slack. Conversely, the increase of government spending is higher under coordination, because policymakers anticipate that

[3]See Corsetti and Pesenti (2001), Benigno and Benigno (2003), De Paoli (2009) and Forlati (2015) for different perspectives on the terms-of-trade externality in the context of monetary policy.

[4]We abstract from non-conventional policies such as forward guidance (Eggertsson and Woodford, 2003) or credit policies by the central bank (see, e.g., Cúrdia and Woodford, 2011). These policies are arguably an imperfect substitute for conventional policies, if only because they are not very well understood and hence controversial (see, e.g., Rogoff, 2016; Giannoni et al., 2016).

the terms of trade remain unaffected by a policy response which is common across countries. At the same time, such a response is expected to boost union-wide inflation (rather than an individual country's terms of trade). This is desirable at the ELB, because expected inflation lowers the real interest rate. We illustrate that the fiscal *stimulus gap* due to the lack of coordination can be quantitatively significant.

This result is specific to currency unions, as a comparison with Cook and Devereux (2011) makes clear. Their analysis characterizes optimal cooperative fiscal policies in a two-country model while assuming that the exchange rate is flexible. Cook and Devereux find that, even as the ELB binds, "there is little case for coordinated global fiscal expansion." In their environment a fiscal expansion in one country generates negative output spillovers if the ELB binds and, in sharp contrast to our results, *depreciates* the domestic terms of trade. Intuitively, higher domestic demand raises inflation and lowers real interest rates in the domestic economy if monetary policy is constrained by the ELB. The nominal exchange rate, if free to adjust, depreciates. This, in turn, more than offsets the effect of higher domestic prices on the terms of trade and reduces the need for coordination.[5]

Our main focus is on discretionary policies, because if policymakers are able to commit, monetary policy is well equipped to stabilize the economy even if it is temporarily constrained by the ELB (Eggertsson and Woodford, 2003). As a result, the ELB is less of an issue and the benefits from fiscal stabilization are small, as Schmidt (2013) shows in a closed-economy context.[6] However, as a benchmark we also compute the optimal fiscal policy response assuming both coordination and commitment.[7] In this case, because we consider a common shock in a perfectly symmetric currency union, the solution is identical to what one would obtain for a closed-economy model. For the case of discretion, we find that there are sizeable benefits from coordinating the fiscal response to the crisis. Moreover, in our analysis these benefits are not offset by credibility problems which may arise under discretion and can be amplified under international policy coordination (Rogoff, 1985). This is because we

[5] Cook and Devereux (2013) discuss in detail the "perverse response" of relative prices at the ELB. Corsetti et al. (2017) offer a systematic comparison of the terms of trade response to government spending in a small open economy: higher spending appreciates the terms of trade under a peg, but depreciates them under flexible exchange rates if monetary policy is constrained by the ELB. However, they do not study optimal fiscal policy.

[6] In principle one may assume that monetary policy is unable to commit, but that fiscal policy is able to do so. As Werning (2012) notes, because of decision and implementation lags, there is an aspect of inherent commitment to fiscal policy. It might be of limited practical relevance, however, because supplements to the budget or mid-year budgets are frequently implemented (Perotti, 2004).

[7] We compute the solution on the basis of the algorithm put forward by Eggertsson and Woodford (2003).

study policy responses around a steady state from which discretionary policymakers have no temptation to deviate. Within our framework we also recoup two results which have already been established in the literature, but are crucial to put our main result into perspective. First, we confirm an earlier finding of Turnovsky (1988) and Devereux (1991): absent coordination policymakers choose *too high* a level of government spending in steady state. This is because governments seek to improve (that is, appreciate) their country's terms of trade through purchases of domestically produced goods.[8] Hence, in steady state the terms-of-trade externality has the opposite effect than in the crisis scenario, because stronger terms of trade are beneficial in the long run, as the economy operates at full capacity.

Second, we also contrast government spending multipliers, that is, the percentage change of domestic output given a (possibly non-optimal) increase of government spending by one percent of GDP in the entire union and in the domestic economy only. In line with earlier work by Fahri and Werning (2016), we find that the multiplier is larger than unity in the first case, provided the ELB binds, but smaller than unity in the second case. This result obtains because a unilateral increase of government spending appreciates the terms of trade and thus crowds out private expenditure. Instead, the cooperative policy, common to all countries, raises expected inflation at the union level, thus crowding-in private expenditure at the ELB.[9]

Our analysis also relates to a number of other recent studies. Blanchard et al. (2016) calibrate a two-country model to capture key features of the euro area, notably of its core and periphery. They show through model simulations that increasing government spending in the core generates significant welfare gains. Evers (2015) studies the performance of alternative fiscal arrangements in a quantitative model of a currency union. He finds that a centralized fiscal authority dominates a regime of fiscal transfers as well as a regime of decentralized fiscal decision making. Other work has focused on the coordination of

[8] Epifani and Gancia (2009) find that this mechanism may account for the size of the public sector in open economies. In particular, their findings suggest that the terms-of-trade externality rather than a demand for insurance causes the public sector to grow with trade openness. Chari and Kehoe (1990) stress that the behavior of cooperative and non-cooperative fiscal policies converge as countries' market power goes to zero. In our setup countries retain market power even though we consider a continuum of small economies.

[9] An alternative perspective emphasizes monetary conditions: at the country level there is a de facto target for the price level, given by purchasing power parity. Any inflationary impulse due to fiscal policy thus triggers an offsetting deflationary tendency and causes the long term real interest rate to rise on impact (Corsetti et al., 2013a). At the union level, absent a price level target, the inflationary impulse due to higher government spending reduces real interest rates at the ELB.

debt and deficit policies in currency unions (Beetsma and Uhlig, 1999; Krogstrup and Wyplosz, 2010). We abstract from this aspect, as Ricardian equivalence obtains in our model. Moreover, we stress that our analysis disregards complications due to sovereign risk. However, both aspects are likely to further strengthen the case for coordination in currency unions stuck at the ELB (Corsetti et al., 2014).

The remainder of the paper is structured as follows. In Section 1.2 we describe the basic setup of the model. It also contrasts government spending multipliers at the union and the country level, once the ELB binds. In Section 1.3 we analyze the need for coordination by determining optimal government spending with and without coordination. Section 1.4 provides a quantitative assessment. Section 1.5 concludes.

1.2 Model

Our analysis is based on the model of Galí and Monacelli (2008). There is a currency union which consists of a continuum of countries, each a small open economy indexed by $i \in [0, 1]$. Each economy features a representative household, a continuum of monopolistically competitive firms and a fiscal authority. Monetary policy is conducted at the union level. We consider two dimensions which are absent in Galí and Monacelli (2008). First, we allow for the possibility that the ELB constrains monetary policy because of a union-wide contractionary shock. Second, we compute optimal fiscal policies when there is no coordination across countries.[10] Our exposition focuses on the model structure in terms of preferences and technology. In a second step, we state the linearized equilibrium conditions at the country and the union level. Readers may consult Galí and Monacelli (2008) for further details on the derivations.

1.2.1 Model structure

In what follows we briefly outline the problem of households, the fiscal authority, firms and monetary policy.

[10] Forlati (2009) also analyzes optimal fiscal policy in the absence of coordination within the Galí-Monacelli model. Her focus is on the interaction of monetary and fiscal policy without considering an ELB.

Households

A representative household in country i has preferences over private consumption, C_t^i, public consumption, G_t^i, and labor, N_t^i, given by

$$U(C_t^i, N_t^i, G_t^i) = (1-\chi)\log C_t^i + \chi \log G_t^i - \frac{\left(N_t^i\right)^{1+\varphi}}{1+\varphi},$$

where parameter $\chi \in (0,1)$ determines the relative weights of private and public consumption. $\varphi > 0$ is the inverse of the Frisch elasticity of labor supply. Private consumption is a composite of domestically produced goods, $C_{i,t}^i$, and imported goods, $C_{F,t}^i$:

$$C_t^i \equiv \frac{\left(C_{i,t}^i\right)^{1-\alpha}\left(C_{F,t}^i\right)^{\alpha}}{(1-\alpha)^{1-\alpha}\,\alpha^{\alpha}}.$$

Parameter $\alpha \in (0,1)$ measures the openness of the economy. Because country i has zero weight in the union, $\alpha < 1$ implies that there is home bias in consumption which accounts for deviations from purchasing power parity in the short run. Domestically produced goods are a CES basket of product varieties:

$$C_{i,t}^i \equiv \left(\int_0^1 C_{i,t}^i(j)^{\frac{\varepsilon-1}{\varepsilon}}dj\right)^{\frac{\varepsilon}{\varepsilon-1}}, \text{ with } \varepsilon > 1. \tag{1.1}$$

Here $C_{i,t}^i(j)$ denotes country i's consumption of variety $j \in [0,1]$ produced in country i. Parameter $\varepsilon > 1$ denotes the elasticity of substitution between different varieties of goods produced within each country. Consumption of imported goods, in turn, is defined as follows:

$$C_{F,t}^i \equiv \exp\int_0^1 c_{f,t}^i df,$$

with $c_{f,t}^i \equiv \log C_{f,t}^i$ and $f \in [0,1]$. The index $C_{f,t}^i$ is defined analogously to (1.1), with an appropriate normalization (Galí and Monacelli, 2015).

Given the definitions above, minimizing expenditures gives rise to demand functions for product varieties. For instance, domestic demand for generic good j is given by

$$C_{i,t}^i(j) = \left(\frac{P_t^i(j)}{P_t^i}\right)^{-\varepsilon} C_{i,t}^i,$$

where $P_t^i(j)$ is the price of good j and $P_t^i \equiv \left(\int_0^1 P_t^i(j)^{1-\epsilon} dj\right)^{\frac{1}{1-\epsilon}}$ is the domestic (producer) price index. Country-i demand for a generic country-f good j, that is $C_{f,t}^i(j)$, is given analogously as well as the producer price index in country f, that is P_t^f.

The optimal allocation between domestic and foreign goods requires

$$C_{i,t}^i = (1-\alpha)\left(\frac{P_t^i}{P_{c,t}^i}\right)^{-1} C_t^i, \qquad C_{F,t}^i = \alpha\left(\frac{P_t^*}{P_{c,t}^i}\right)^{-1} C_t^i,$$

where $P_t^* \equiv \exp \int_0^1 p_t^f df$ is the union-wide price index with $p_t^f \equiv \log P_t^f$. The consumer price index (CPI) is given by $P_{c,t}^i \equiv (P_t^i)^{1-\alpha}(P_t^*)^{\alpha}$. In the following we focus on the producer price index, P_t^i, which is related to the CPI according to $P_t^i = P_{c,t}^i(S_t^i)^{\alpha}$, where $S_t^i \equiv P_t^*/P_t^i$ denotes the terms of trade.

Households trade a complete set of state-contingent securities which provides insurance against country-specific shocks.[11] They maximize expected discounted lifetime utility subject to a sequence of budget constraints:

$$\begin{aligned} &\max_{\{C_t^i, N_t^i, A_t^i\}_{t=0}^{\infty}} E_0 \sum_{t=0}^{\infty} \beta^t U(C_t^i, N_t^i, G_t^i) \\ &\quad \text{s.t.} \quad P_{c,t}^i C_t^i + E_t\{Q_{t,t+1} A_{t+1}^i\} \leq A_t^i + W_t^i N_t^i + \mathcal{P}_t^i - T_t^i. \end{aligned}$$

where A_t^i denotes the portfolio of nominal assets and $Q_{t,t+1}$ is the nominal stochastic discount factor (common across countries). Ponzi schemes are not permitted. W_t^i is the nominal wage and $\mathcal{P}_t^i$ are firm profits, rebated to households in a lump-sum fashion. T_t^i are lump-sum taxes. We consider the case of distortinary taxes in Section 1.4.3 below. Parameter $\beta \in (0,1)$ is the subjective discount factor.

[11] For instance, idiosyncratic technology shocks as in Galí and Monacelli (2008). As we analyze optimal policy in response to an aggregate shock that pushes the currency union at the ELB we abstract from country-specific shocks in our analysis.

Fiscal authority

Public consumption is composed of domestically produced goods as in (1.1) and the fiscal authority allocates expenditures in a cost minimizing manner. The resulting demand function for a generic good j is given by:

$$G_t^i(j) = \left(\frac{P_t^i(j)}{P_t^i} \right)^{-\varepsilon} G_t^i.$$

Aggregate expenditure, G_t^i, remains to be determined below. Taxes adjust to balance the budget in each period:

$$T_t^i = P_t^i G_t^i + \tau^i W_t^i N_t^i. \tag{1.2}$$

where τ^i is a (constant) employment subsidy paid to domestic firms. If set appropriately it ensures the efficiency of the steady state under monopolistic competition.

Firms

In each country, there is a continuum of monopolistically competitive firms, each of which produces a differentiated good $Y_t^i(j)$. These goods are traded across countries and the law of one price is assumed to hold. Firms cannot adjust their price $P_t^i(j)$ every period. Instead, as in Calvo (1983), they may reset prices in a given period with probability $1-\theta$, while their current price remains in effect with probability $\theta \in (0,1)$. The probability of resetting the price is independent of the last adjustment. Firms hire labor $N_t^i(j)$ and produce with a linear technology $Y_t^i(j) = N_t^i(j)$ in order to satisfy the level of demand at a given price. The objective of a generic firm $j \in [0,1]$ is to maximize discounted, expected nominal payoffs taking the demand for its product into account. The optimization problem is given by:

$$\max_{\bar{P}_t^i(j)} \sum_{k=0}^{\infty} \theta^k E_t \left\{ Q_{t,t+k} Y_{t+k}^i(j) (\bar{P}_t^i(j) - (1-\tau^i) W_{t+k}^i) \right\}$$
$$\text{s.t.} \quad Y_{t+k}^i(j) = \left(\frac{\bar{P}_t^i(j)}{P_{t+k}^i} \right)^{-\varepsilon} Y_{t+k}^i,$$

where $\bar{P}_t^i(j)$ is the optimal price, set in period t.

Monetary policy

Monetary policy is conducted at the union level. The policy instrument is the nominal interest rate, that is, the yield on a nominally riskless one-period discount bond: $1 + i_t^* \equiv \frac{1}{E_t\{Q_{t,t+1}\}}$. The objective of monetary policy is to maintain price stability, that is, zero inflation at the union level.[12] Importantly, monetary policy may be constrained by an ELB. Specifically, in what follows we assume that $i_t^* \geq 0$, that is, we assume the effective lower bound to be zero. While the actual lower bound is arguably somewhat below zero, this is of little consequence in the context of our analysis. Below we specify an interest-rate rule which implements price stability subject to the ELB constraint.

1.2.2 Equilibrium conditions for approximate model

We consider a log-linear approximation to the optimality and market-clearing conditions around a symmetric, zero-inflation steady state. We use hats to denote log-deviations of a variable from its steady-state value. For a generic variable X_t we define $x_t \equiv \log X_t$ and $\hat{x}_t = \log(X_t/X)$. Union-wide variables are obtained by integrating over all countries in the union: $\hat{x}_t^* = \int_0^1 \hat{x}_t^i di$.

First, goods-market clearing and integrating over all goods gives for country i

$$\hat{y}_t^i = (1-\gamma)(\hat{c}_t^i + \alpha s_t^i) + \gamma \hat{g}_t^i. \tag{1.3}$$

Parameter γ denotes the steady-state ratio of government consumption to output. The above equation links domestic output $\hat{y}_t^i$ to domestic consumption $\hat{c}_t^i$, the terms of trade s_t^i and domestic government spending $\hat{g}_t^i$. Further, the assumption of complete markets gives rise to the following risk sharing condition:

$$\hat{c}_t^i = \hat{c}_t^* + (1-\alpha)s_t^i. \tag{1.4}$$

Combining it with (1.3) gives

$$\hat{y}_t^i = \gamma \hat{g}_t^i + (1-\gamma)\hat{c}_t^* + (1-\gamma)s_t^i. \tag{1.5}$$

Integrating equation (1.5) over all countries $i \in [0,1]$ and noting that $\int_0^1 s_t^i di = 0$ leads to

[12] In the context of our model this is the optimal discretionary policy under fiscal coordination.

the union-wide market clearing condition

$$\hat{y}_t^* = \gamma \hat{g}_t^* + (1-\gamma)\hat{c}_t^*. \tag{1.6}$$

Combining (1.5) and (1.6), we can rewrite market clearing at the country level as follows:

$$\hat{y}_t^i - \hat{y}_t^* = \gamma(\hat{g}_t^i - \hat{g}_t^*) + (1-\gamma)s_t^i. \tag{1.7}$$

Integrating country-specific Euler equations over all countries $i \in [0,1]$ and combining it with (1.6) yields a union-wide dynamic IS curve:

$$\hat{y}_t^* = E_t\{\hat{y}_{t+1}^*\} - (1-\gamma)(i_t^* - E_t\{\pi_{t+1}^*\} - r_t) - \gamma E_t\{\hat{g}_{t+1}^*\} + \gamma \hat{g}_t^* \tag{1.8}$$

with $r_t \equiv -\log\beta - \Delta_t$. As in Woodford (2011), Δ_t denotes a spread between the interest rate set by the central bank and the one relevant for private sector decisions. It reflects frictions in financial intermediation which we do not model explicitly, but permit to vary exogenously.[13] If this spread becomes large enough, monetary policy becomes constrained by the ELB. In what follows, we assume that monetary policy follows a Taylor rule unless it is constrained by the ELB. Specifically, we posit the following:

$$i_t^* = \max\left\{r_t + \phi_\pi \pi_t^*, 0\right\}. \tag{1.9}$$

We restrict $\phi_\pi > 1$. By following this rule monetary policy fully stabilizes inflation and output at the union level (as long as $\hat{g}_t^* = 0$), unless it is constrained by the ELB.

Optimal price-setting behavior of firms implies the following variant of the New Keynesian Phillips curve:

$$\pi_t^i = \beta E_t\{\pi_{t+1}^i\} + \lambda\left(\frac{1}{1-\gamma} + \varphi\right)\hat{y}_t^i - \frac{\lambda\gamma}{1-\gamma}\hat{g}_t^i, \tag{1.10}$$

with $\lambda \equiv \frac{(1-\beta\theta)(1-\theta)}{\theta}$ and where $\pi_t^i = p_t^i - p_{t-1}^i$ denotes the inflation rate. Integrating

[13]Cúrdia and Woodford (2016) provide a microfoundation in a model which accounts for household heterogeneity and borrowing and lending across households.

equation (1.10) over all countries $i \in [0,1]$ gives the union-wide Phillips curve:

$$\pi_t^* = \beta E_t\{\pi_{t+1}^*\} + \lambda\left(\frac{1}{1-\gamma}+\varphi\right)\hat{y}_t^* - \frac{\lambda\gamma}{1-\gamma}\hat{g}_t^*. \tag{1.11}$$

From the definition of the terms of trade it follows that

$$\pi_t^i = \pi_t^* - s_t^i + s_{t-1}^i. \tag{1.12}$$

Further, we note that equilibrium conditions (1.7) and (1.10)-(1.12) imply the following second order stochastic difference equation for the terms of trade (see Galí and Monacelli, 2005b)

$$s_t^i = \omega s_{t-1}^i + \omega\beta E_t\{s_{t+1}^i\} - \omega\lambda\varphi\gamma(\hat{g}_t^i - \hat{g}_t^*),$$

where $\omega \equiv \frac{1}{1+\beta+\lambda[1+\varphi(1-\gamma)]} \in [0, \frac{1}{1+\beta})$. The above equation has a unique stable solution

$$s_t^i = \delta s_{t-1}^i + \delta\lambda\varphi\gamma\sum_{k=0}^{\infty}(\beta\delta)^k E_t\{\hat{g}_{t+k}^* - \hat{g}_{t+k}^i\}, \tag{1.13}$$

with $\delta \equiv \frac{1-\sqrt{1-4\beta\omega^2}}{2\omega\beta} \in (0,1)$.

Definition of equilibrium. *Given initial conditions (s_{-1}) as well as $\{\Delta_t\}_{t=0}^{\infty}$ an equilibrium is a collection of*

1. *country-specific stochastic processes $\{\hat{y}_t^i, \pi_t^i, s_t^i\}_{t=0}^{\infty}$ for all $i \in [0,1]$*
2. *union-wide stochastic processes $\{\hat{y}_t^*, \pi_t^*\}_{t=0}^{\infty}$ with $\hat{y}_t^* = \int_0^1 \hat{y}_t^i di$ and $\pi_t^* = \int_0^1 \pi_t^i di$*

such that for given $\{\hat{g}_t^i\}_{t=0}^{\infty}$ for all $i \in [0,1]$ with $\hat{g}_t^* = \int_0^1 \hat{g}_t^i di$ and the path for the nominal interest rate $\{i_t^*\}_{t=0}^{\infty}$ determined by (1.9)

3. *equilibrium conditions* (1.7), (1.10) *and* (1.12) *are satisfied for each country i and*
4. *equilibrium conditions* (1.8) *and* (1.11) *are satisfied on the union level.*

Effective-lower-bound (or "crisis") scenario. In our analysis below, we consider a scenario where the ELB binds because the interest rate spread increases temporarily.

Specifically, as in Woodford (2011), we assume a Markov structure for Δ_t. It rises temporarily to a value Δ_L such that $r_L < 0$. The shock remains operative with probability μ and is sufficiently large for the ELB to become binding. Once the shock disappears there are no more future ELB episodes.[14] Formally, equation (1.9) implies that $i_t^* = 0$ for as long as the shock lasts, independently on the conduct of fiscal policy.[15] With probability $1-\mu$ the spread disappears (and thus the whole economy returns permanently to the steady state). Moreover, defining $\kappa \equiv \lambda\left(\frac{1}{1-\gamma} + \varphi\right)$, we impose the parametric restriction $(1-\mu)(1-\beta\mu) > (1-\gamma)\mu\kappa$ for the equilibrium to be uniquely determined (Woodford, 2011).

1.2.3 Impact multipliers: union-wide vs country-specific fiscal impulse

In this section, to set the stage for our main results in Section 1.3, we solve for the government spending multiplier on output. That is, we determine by how much country-specific output changes, given an increase of government consumption by one percent of output. Our focus is on how the multiplier differs depending on whether there is a union-wide or a country-specific variation of government consumption. As a union-wide fiscal impulse impacts the individual countries symmetrically, this scenario is equivalent to the closed-economy setting in Woodford (2011). Instead, a country-specific fiscal impulse impacts domestic output directly, but also indirectly via the terms of trade. This scenario is thus equivalent to the small-open-economy settings in Corsetti et al. (2013a) and Fahri and Werning (2016). We briefly revisit their results within our framework.

Consider first the union-wide fiscal impulse in the ELB scenario. We assume that government spending is increased in every country by the same amount as long as the ELB remains binding. In this case, given the assumptions spelled out above, union-wide variables take a constant value x_L^*, as long as the shock persists und the union-wide Phillips curve

[14] Assuming such an absorbing state ensures the tractability of the model and allows us to derive closed-form results for discretionary policies, because in this case the economy only visits two states: the ELB state and the steady state. Adam and Billi (2006, 2007) compute optimal (monetary) policy while allowing the ELB to bind occasionally.

[15] Schmidt (2013) and Erceg and Lindé (2014) consider endogenous exit from the ELB due to fiscal-policy measures. We assume instead that the decline of r_t is sufficiently large for the ELB to remain binding also in the presence of optimal fiscal stabilization.

and the IS equation simplify to

$$\pi_L^* = \frac{1}{1-\beta\mu}\kappa\left(\hat{y}_L^* - \frac{\bar{\sigma}\gamma}{\bar{\sigma}+\varphi}\hat{g}_L^*\right), \tag{1.14}$$

$$(1-\mu)(\hat{y}_L^* - \gamma\hat{g}_L^*) = (1-\gamma)\mu\pi_L^* + (1-\gamma)r_L, \tag{1.15}$$

with $\bar{\sigma} \equiv \frac{1}{1-\gamma}$.

We solve the above system for $\hat{y}_L^*$ as a function of r_L and $\hat{g}_L^*$. This gives:

$$\hat{y}_L^* = \frac{(1-\gamma)(1-\beta\mu)}{(1-\mu)(1-\beta\mu)-(1-\gamma)\mu\kappa}r_L + \frac{(1-\mu)(1-\beta\mu)\gamma - (1-\gamma)\mu\kappa\frac{\gamma\bar{\sigma}}{\bar{\sigma}+\varphi}}{(1-\mu)(1-\beta\mu)-(1-\gamma)\mu\kappa}\hat{g}_L^*. \tag{1.16}$$

In order to determine the multiplier, we divide the derivative of $\hat{y}_L^*$ with respect to $\hat{g}_L^*$ by the steady-state share of government spending, γ:

$$\frac{1}{\gamma}\frac{\partial\hat{y}_L^*}{\partial\hat{g}_L^*} = \frac{(1-\mu)(1-\beta\mu)-(1-\gamma)\mu\kappa\frac{\bar{\sigma}}{\bar{\sigma}+\varphi}}{(1-\mu)(1-\beta\mu)-(1-\gamma)\mu\kappa} \geq 1.$$

At the union level, we thus find that the multiplier is bounded from below by unity (Woodford, 2011). Intuitively, higher government spending reduces real interest rates at the ELB, because the expected inflationary impact of higher spending is not matched by higher nominal interest rates. Hence, private-sector spending is crowded in.

We now turn to the effect of a country-specific fiscal impulse. In this case, we set union-wide variables to zero and to ensure comparability with the union-wide fiscal impulse we assume that government spending in country i follows a two-state Markov switching process. Initially, government spending exceeds its steady state level $\hat{g}_L^i > 0$; it does so with probability μ in the next period too and returns to steady state with probability $1-\mu$.

Specifically, equations (1.7) and (1.13), evaluated in the impact period of the spending increase read as follows

$$\hat{y}_1^i = \gamma\hat{g}_L^i - (1-\gamma)p_1^i$$
$$p_1^i = \frac{\delta\lambda\varphi\gamma}{1-\beta\delta\mu}\hat{g}_L^i.$$

Combining both equations, we obtain the government spending multiplier in the impact period.[16] It is given by

$$\frac{1}{\gamma}\frac{\partial \hat{y}_1^i}{\partial \hat{g}_L^i} = 1 - (1-\gamma)\underbrace{\frac{\delta\lambda\varphi}{1-\beta\delta\mu}}_{\geq 0} \leq 1.$$

The upper bound of unity is reached when prices are completely sticky ($\lambda \to 0$). To the extent that prices are somewhat flexible, private-sector spending at the country level is crowded out by higher government consumption. Its inflationary impact appreciates the terms of trade which, in turn, calls for reduced consumption in country i, see equation (1.4). Equivalently, (relative) purchasing power parity requires that the price level reverts back to its pre-shock level in the long run. Given unchanged nominal interest rates in the currency union, future deflation induces long-term real interest rates to rise on impact. Still, the crowding-out effect of a country-specific stimulus in a currency union is limited relative to when the country operates a flexible exchange rate system. In other words, the multiplier is larger under a fixed exchange rate than under flexible exchange rates (see, for further discussion and evidence, Corsetti et al., 2013a; Born et al., 2013).

Taken together, we obtain the following ranking of the government spending multiplier on country-specific output, considering a union-wide and country-specific spending increase, respectively:

$$\frac{1}{\gamma}\frac{d\hat{y}_1^i}{d\hat{g}_L^i} \leq 1 \leq \frac{1}{\gamma}\frac{d\hat{y}_L^*}{d\hat{g}_L^*}.$$

Fahri and Werning (2016) obtain this result as a closed-form solution of the continuous-time version of the New Keynesian model.[17]

[16] Because a country-specific fiscal impulse impacts the terms of trade, the output effect of government spending changes over time even though the size of the impulse does not. We focus on the impact effect.

[17] Erceg and Lindé (2012) also compute spending multipliers for a small open economy. Assuming an exchange rate peg, they show that multipliers are always below unity. For the case of flexible exchange rates, they stress that at the ELB the multiplier exceeds unity only if prices are sufficiently flexible. To account for this finding note that, in contrast to us, they do not assume that government spending is raised only for as long as the ELB binds. Nakamura and Steinsson (2014), in turn, show that multipliers are high within a currency union when compared to the multiplier at the union level in the absence of a binding ELB constraint. Acconcia et al. (2014) find for Italian data that variations in local government spending have fairly strong output effects.

1.3 Optimal policy

We now turn to optimal fiscal policy. In particular, we distinguish between a scenario of coordination and one without, both with regards to the steady state and to the ELB scenario. In each case, policymakers chose government consumption in order to maximize household welfare.

1.3.1 Steady state

We consider optimal fiscal policy in the steady state first. We compute the symmetric steady state as the solution to the social planner problem and discuss how it can be decentralized.

Under coordination the social planner (of the union) maximizes union-wide welfare subject to the production function and the goods-market-clearing condition. Formally, we have

$$\max \int_0^1 \left((1-\chi) \log C^i + \chi \log G^i - \frac{\left(N^i\right)^{1+\varphi}}{1+\varphi} \right) di \tag{1.17}$$

$$\text{s.t.} \quad Y^i = N^i$$

$$Y^i = C_i^i + \int_0^1 C_i^f \, df + G^i,$$

for all $i \in [0,1]$ where C_i^f denotes country f's consumption of goods produced in country i. In addition, optimality requires that varieties are produced and consumed in equal quantities in each country (which is already assumed in the above constraints). Solving the planner problem gives rise to the following steady-state relations (for each country $i \in [0,1]$), see Galí and Monacelli (2008):

$$\gamma^C \equiv \left(\frac{G}{Y}\right)^C = \chi; \quad Y^C = 1, \tag{1.18}$$

where index C refers to the case of coordination. The social planner solution can be decentralized as a symmetric, zero-inflation steady state by letting the government provide public goods according to (1.18) and by setting the labor subsidy to offset distortions due to monopolistic competition:

$$\tau^C = \frac{1}{\varepsilon}.$$

We now turn to the case without coordination. Here, the social planner (in a given country i) maximizes domestic welfare only, subject to the production function, the market-clearing condition and the risk-sharing condition. In this regard, the planner in country i is subject to the same constraints vis-à-vis the rest of the union as is country i in a decentralized equilibrium (Galí and Monacelli, 2005a). The planner also takes consumption in the rest of the union C^* as given. Formally, we have

$$\begin{aligned} &\max\,(1-\chi)\log C^i + \chi \log G^i - \frac{\left(N^i\right)^{1+\varphi}}{1+\varphi} \qquad (1.19)\\ &\text{s.t.} \qquad Y^i = N^i \\ &\qquad\qquad Y^i = C^i(S^i)^\alpha + G^i \\ &\qquad\qquad C^i = C^*(S^i)^{1-\alpha}. \end{aligned}$$

Again, in an optimal allocation varieties are produced and consumed in equal quantities in a given country. In a symmetric Nash equilibrium, optimality requires the following to hold in every country $i \in [0,1]$, see Appendix 1.A.1:

$$\begin{aligned} \gamma^N \equiv \left(\frac{G}{Y}\right)^N &= \frac{\chi}{(1-\alpha)(1-\chi)+\chi} \in (0,1) \qquad (1.20)\\ Y^N &= \left[(1-\alpha)(1-\chi)+\chi\right]^{\frac{1}{1+\varphi}}, \end{aligned}$$

where index N refers to the case of *no* coordination. Comparing the outcome under coordination and without, we observe that the government-consumption-to-output ratio is higher in the latter case: $\gamma^N > \gamma^C$. Furthermore, the level of government spending without coordination exceeds the level under coordination: $G^N > G^C$, even though output is lower $Y^N < Y^C$, see also Appendix 1.A.1.

This confirms earlier findings by Turnovsky (1988) and Devereux (1991) according to which government consumption without coordination accounts for an excessively large share of output. Intuitively, each government tries to improve the domestic terms of trade by increasing domestic demand. In a symmetric Nash equilibrium, however, the terms of trade are equal to unity. Government consumption is higher and output is lower relative to the case of coordination. Because of the terms-of-trade externality, the steady state in the absence of coordination is inefficient: welfare is lower than in case of coordination.

The social planner solution in the absence of coordination can be decentralized as a symmetric, zero-inflation steady state by letting the government provide public goods according to (1.20) and by choosing the following labor subsidy, see Appendix 1.A.2:

$$1-\tau^N = \left(1-\frac{1}{\varepsilon}\right)(1-\alpha)^{-1}.$$

Here, as explained by Galí and Monacelli (2005a), the optimal employment subsidy offsets the combined effects of market power and the terms of trade distortion such that the flexible price equilibrium allocation is optimal (from the viewpoint of the social planner in the absence of coordination). Hence, in our analysis below, there is no average inflation (or deflation) bias. Finally, zero inflation ensures that the same amount of each variety is produced and consumed.

1.3.2 Effective lower bound

In order to determine the optimal discretionary fiscal policy at the ELB, we pursue a linear-quadratic approach. First, we approximate household welfare up to second order. In the case of coordination we approximate the welfare function as well as the equilibrium conditions around the steady state with coordination. Instead, in the case without coordination we approximate around the steady state without coordination. In fact, Galí and Monacelli (2008) already provide an approximation of household welfare for the case of coordination. We provide details on the derivation in the absence of coordination in Appendix 1.B.[18] Second, we determine the optimal discretionary fiscal policy as we maximize the welfare functions subject to the equilibrium conditions, once approximated around the steady state with coordination and once without.[19]

Consider first the case of coordination. Here we focus on the symmetric solution, that is, $x_t^i = x_t^*$ for all $i \in [0,1]$, because we analyze the effects of a union-wide shock and

[18] In the absence of coordination there are linear terms in a second-order approximation to household utility. We follow the approach of Benigno and Woodford (2006) and substitute for these terms using a second-order approximation to the market-clearing condition.

[19] Clarida et al. (2002) also compare optimal (monetary) policy under coordination and without, each evaluated around the respective steady state.

assume that countries are identical. Under coordination the single policymaker maximizes union-wide welfare by choosing government consumption in a discretionary way subject to the New Keynesian Phillips curve, (1.11), and an inequality constraint which consolidates the dynamic IS equation, (1.8) and the interest-rate rule (1.9). Hence, optimization is subject to the ELB. Formally, assuming discretionary policy making, the optimization problem is given by

$$\begin{aligned} \max_{\pi_t^*,\hat{y}_t^*,\hat{g}_t^*} \mathbb{W}_t^* &\simeq -\frac{1}{2}\left(\frac{\varepsilon}{\lambda}(\pi_t^*)^2 + (1+\varphi)(\hat{y}_t^*)^2 + \frac{\gamma^C}{1-\gamma^C}(\hat{g}_t^* - \hat{y}_t^*)^2\right) \\ \text{s.t.} \quad \pi_t^* &= \lambda\left(\frac{1}{1-\gamma^C}+\varphi\right)\hat{y}_t^* - \frac{\lambda\gamma^C}{1-\gamma^C}\hat{g}_t^* + \nu_{0,t}^* \\ \hat{y}_t^* &\leq \gamma^C \hat{g}_t^* + \nu_{1,t}^*, \end{aligned} \tag{1.21}$$

where $\nu_{0,t}^*$ and $\nu_{1,t}^*$ collect expectation terms which are beyond the control of the policy maker under discretion. In the expression above the welfare function features in addition to inflation and the output term a fiscal gap, that is, the deviation of the spending-to-output ratio from steady state. In Appendix 1.C we show that the solution to (1.21) requires $\pi_t^* = \hat{y}_t^* = \hat{g}_t^* = 0$ as long as the monetary authority is not constrained by the ELB. Hence, in this case the economy is perfectly stabilized at the steady state and the government-consumption-to-output ratio is at its efficient level. When the ELB is binding, however, $\pi_t^* = \hat{y}_t^* = \hat{g}_t^* = 0$ is no longer feasible. In that case we find that optimal government spending is characterized by the following condition:

$$\pi_t^{*,C} + \frac{1}{\varepsilon}\hat{y}_t^{*,C} = -\psi_g^C\, \hat{g}_t^{*,C}, \tag{1.22}$$

where $\psi_g^C \equiv \frac{1}{\varepsilon\varphi} > 0$, see case 2 in Appendix 1.C. Intuitively, as long as output and inflation drop in the crisis scenario (left-hand-side in the expression above), the optimal policy is to raise government spending. The following proposition states the solution for optimal government spending.

Proposition 1. Given the effective-lower-bound scenario under consideration (see Section 1.2.2), the optimal response of government spending under coordination, measured in percentage deviations from the steady state with coordination, $\hat{g}_L^{*,C}$, is given by:

$$\hat{g}_L^{*,C} = -\Theta^C r_L > 0,$$

with

$$\Theta^C \equiv \frac{(1-\gamma^C)\left(\kappa^C\varepsilon + (1-\beta\mu)\right)}{\varepsilon\psi_g^C\Gamma_1^C + (1-\mu)\gamma^C\lambda\varphi\varepsilon + \gamma^C\Gamma_2^C} > 0,$$

where $\kappa^C \equiv \lambda\left(\frac{1}{1-\gamma^C}+\varphi\right)$, $\Gamma_1^C \equiv (1-\mu)(1-\beta\mu)-(1-\gamma^C)\mu\kappa^C$ and $\Gamma_2^C \equiv (1-\mu)(1-\beta\mu)-\mu\lambda$ and $\Gamma_1^C, \Gamma_2^C > 0$ because of the assumption of a uniquely determined equilibrium.

Proof. See Appendix 1.D. ■

Hence, we find that it is optimal to raise government spending under coordination, once the ELB binds. The increase of government spending is time-invariant in the crisis scenario because of the Markov structure of the shock. Once the ELB ceases to bind, the economy returns to steady state. Our result is in line with Woodford (2011), because in the present context the currency union under coordination is isomorphic to his closed-economy model.

We now turn to optimal government spending in the absence of coordination. In this case, policy choices may differ across countries from an ex-ante perspective and, hence, are expected to impact a country's terms of trade. Given price stickiness, the terms of trade adjust sluggishly in a currency union and hence the (lagged) terms of trade are an endogenous state variable.[20] As a result, the policy problem is inherently dynamic—even under discretion. In this case, even though a policymaker may not directly steer private-sector expectations, current policy decisions impact expectations indirectly via endogenous state variables—an effect which is internalized by the policymaker (see, e.g., Svensson, 1997). We further note that, since the local policymaker takes union-wide variables as given, including the nominal interest rate, the union-wide IS curve is not a constraint for the decision maker and neither is the ELB. Instead, optimization is subject to the market-clearing-cum-risk-sharing condition, (1.7), the country-specific New Keynesian Phillips curve, (1.10), and the evolution of the terms of trade, (1.12). Specifically, under discretion

[20] The sluggish adjustment of the terms of trade has been identified as a key determinant of the macroeconomic adjustment mechanism in monetary unions. Benigno (2004) and Pappa (2004) stress how it distorts the adjustment. Groll and Monacelli (2016), in contrast, relate it to the "intrinsic benefits of monetary unions" in the absence of commitment.

the optimization problem is given by

$$
\begin{aligned}
V(s_{t-1}^i, \pi_t^*, \hat{y}_t^*, \hat{g}_t^*) = \max_{\pi_t^i, \hat{y}_t^i, \hat{g}_t^i, s_t^i} \Bigg[& -\frac{1}{2}\left(\frac{\varepsilon}{\lambda}(\pi_t^i)^2 + (1+\varphi)(\hat{y}_t^i)^2 + \frac{\gamma^N}{1-\gamma^N}\left(\hat{g}_t^i - \hat{y}_t^i\right)^2\right) \\
& + \beta E_t V(s_t^i, \pi_{t+1}^*, \hat{y}_{t+1}^*, \hat{g}_{t+1}^*)\Bigg] \qquad (1.23)
\end{aligned}
$$

$$
\begin{aligned}
\text{s.t.} \quad \hat{y}_t^i =& \hat{y}_t^* + \gamma^N(\hat{g}_t^i - \hat{g}_t^*) + (1-\gamma^N)s_t^i \\
\pi_t^i =& \beta E_t\{\pi_{t+1}^i\} + \lambda\left(\frac{1}{1-\gamma^N} + \varphi\right)\hat{y}_t^i - \frac{\lambda\gamma^N}{1-\gamma^N}\hat{g}_t^i \\
\pi_t^i =& \pi_t^* - s_t^i + s_{t-1}^i \\
& \text{and } E_t\{\pi_{t+1}^i\} \text{ given.}
\end{aligned}
$$

In the expression above V is the value function. The solution to (1.23) requires the following (consolidated) first-order condition to be satisfied (see Appendix 1.E):

$$
-\lambda\beta E_t \frac{\partial V(s_t^{i,N}, \pi_{t+1}^{*,N}, \hat{y}_{t+1}^{*,N}, \hat{g}_{t+1}^{*,N})}{\partial s_t^{i,N}} + \beta\left(\hat{y}_t^{i,N} + \frac{1}{\varphi}\hat{g}_t^{i,N}\right)\frac{\partial E_t\{\pi_{t+1}^{i,N}\}}{\partial s_t^{i,N}} + \pi_t^{i,N} + \frac{1}{\varepsilon}\hat{y}_t^{i,N} = -\psi_g^N \hat{g}_t^{i,N}, \qquad (1.24)
$$

with $\psi_g^N \equiv \frac{1}{\varepsilon\varphi}(\lambda\varphi + (1+\lambda))$.

To develop some intuition, it is instructive to contrast optimality condition (1.24) with the one derived under coordination, equation (1.22). For this purpose we abstract in a first step from the dynamic terms on the left of equation (1.24). We observe that for a given drop of output and inflation in the ELB scenario under consideration, the optimal policy response entails a smaller increase of government spending than in case of coordination, since $\psi_g^N > \psi_g^C$. Intuitively, in the absence of coordination, a local policymaker anticipates that higher government spending appreciates the terms of trade which, in turn, lowers the demand for domestic goods. This effect is absent when government spending is raised simultaneously in all countries under coordination. A non-cooperative policymaker will therefore tend to opt for less fiscal stimulus. The following proposition establishes this formally for the special case which eliminates the dynamic terms in equation (1.24).

Proposition 2. Given the effective-lower-bound scenario (see Section 1.2.2) and assuming a symmetric equilibrium while $\beta \to 0$, the optimal response of government spending w/o coordination, measured in percentage deviations from the steady state w/o coordination,

$\hat{g}_L^{*,N}$, is given by:

$$\hat{g}_L^{*,N} = -\Theta^N_{\beta\to 0}\, r_L > 0,$$

with

$$\Theta^N_{\beta\to 0} \equiv \frac{(1-\gamma^N)\left(\kappa^N\varepsilon + 1\right)}{\varepsilon\psi_g^N\Gamma_1^N + (1-\mu)\gamma^N\lambda\varphi\varepsilon + \gamma^N\Gamma_2^N} > 0,$$

where $\kappa^N \equiv \lambda\left(\frac{1}{1-\gamma^N} + \varphi\right)$, $\Gamma_1^N \equiv (1-\mu) - (1-\gamma^N)\mu\kappa^N$ and $\Gamma_2^N \equiv (1-\mu) - \mu\lambda$ and where $\Gamma_1^N, \Gamma_2^N > 0$ because of the assumption of a uniquely determined equilibrium. Furthermore, it holds that

$$\Theta^N_{\beta\to 0} < \Theta^C_{\beta\to 0},$$

where Θ^C is stated in Proposition 1 above.

Proof. See Appendix 1.F. ■

As in the case of coordination, the optimal response is constant for as long as the ELB binds because in (the symmetric) equilibrium under consideration the terms of trade are unaltered.

In the general case for $\beta \in (0,1)$, the optimal policy also reflects the fact that the terms of trade operate as an endogenous state variable. The first term on the left of equation (1.24) captures the effect that, all else equal, stronger terms of trade (that is, a lower s_t) are expected to persist and to reduce expected future welfare when foreign demand and foreign inflation are weak (as in the ELB scenario). To the extent that higher government spending appreciates the terms of trade, there is thus an additional incentive to opt for less spending in the absence of coordination. This effect reinforces the ordering established in Proposition 2.

Turning to the second term on the left of equation (1.24), note that stronger terms of trade (that is, a lower s_t) reduce inflation expectations, because, as they persist, they will raise the purchasing power of workers and induce downward pressure on wages and inflation (that is, $\partial E_t\{\pi^{i,N}_{t+1}\}/\partial s^{i,N}_t > 0$). Via the Phillips curve, lower expected inflation reduces inflation today. This dynamic terms-of-trade channel attenuates the appreciation of the terms of trade in response to higher government spending (and output) and makes a fiscal stimulus in the absence of cooperation relatively more attractive. Yet this dynamic channel does not overturn the ordering established in Proposition 2 because it merely dampens the appreciation of the terms of trade.

Overall, we thus find that the terms of trade are crucial for optimal policy design, not only in steady state, but also off steady state. Intuitively, without coordination there is excessive government consumption in steady state, because better terms of trade are beneficial in the long run, as the economy operates at full capacity. In the short-run, local policymakers are keen to avoid the terms of trade appreciating too much with higher spending, as it reduces the demand for domestic goods in times of economic slack.

1.4 Quantitative assessment

In this section we explore to which extent our result matters quantitatively as well as the robustness of our results. In this regard we look at alternative parameter values, but we also consider a version of the model with distortionary taxes. Lastly, we compute the level of government spending with and without coordination and the welfare loss which results from the lack of coordination at the ELB.

1.4.1 Baseline

In a first step we define a baseline scenario. For this purpose we assign parameter values by targeting observations for the euro area and the US for the (pre-crisis) period 1999–2006. We treat the US as a benchmark for a currency union in which government spending is set cooperatively. For the euro area, in contrast, we assume that government expenditures are set non-cooperatively.[21]

A period in the model corresponds to one quarter. We set $\chi = 0.148$ in order to match the average share of exhaustive government consumption relative to GDP in the US (see Figure 1.1 above). To match the share of government consumption in the euro area which is equal to 0.196, we set the openness parameter α to 0.2874, see equation (1.20).[22] Further, we set the time-discount factor β to 0.99 and $\theta = 0.925$. Such a high degree of price stickiness appears to be justified in light of the inflation dynamics observed in the context of the crisis (Corsetti et al., 2013b). Moreover, as we illustrate by means of a sensitivity

[21] According to NIPA data 36.3% of exhaustive government expenditures in the US are determined at the federal level. In the EU there is a common budget. However, it is very small and consists mostly of transfers. There is basically no exhaustive government spending administrated at the area-wide level.

[22] The average import share in the euro area during 1999–2006 is closer to 50 percent, but this accounts also for trade with countries outside of the euro area. The share of intra-euro area imports in GDP is about 17 percent according to the Monthly Foreign Trade Statistics of the OECD.

Table 1.1: Parameter values

χ	0.148	Public consumption-GDP ratio
α	0.2874	Import-share in steady state
β	0.99	Discount factor
θ	0.925	Degree of price stickiness
ε	6	Elasticity of substitution
φ	2	Inverse of Frisch elasticity of labor supply
ϕ_π	1.5	Taylor coefficient
r_L	-0.0025	ELB scenario
μ	0.8	Expected duration of ELB

analysis below, understating the extent of price rigidity biases results in favor of fiscal policy coordination. In this sense, a high degree of price rigidity is a conservative choice. We set the elasticity of substitution among varieties to $\varepsilon = 6$. This implies an average markup of 20 percent. We assume $\varphi = 2$, which implies a Frisch elasticity of one half (Chetty et al., 2011). We further assume for monetary policy that $\phi_\pi = 1.5$. In terms of the shock, we assume that $r_L = -0.0025$. This implies a natural rate of interest of -1% (annualized) for the ELB scenario. As a baseline we assume $\mu = 0.8$ which implies that the ELB is expected to bind for 5 quarters. We explore below to what extent results vary with μ, but also explore the robustness of our results with respect to changes in β, θ, φ and α. We also verify that the ELB remains binding as government spending is raised in response to the shock. Table 1 summarizes the parameter values.[23]

Given these parameter values, we solve the model in the absence of coordination numerically as in Soderlind (1999).[24] Appendix 1.G provides details. As we mention above, in the absence of coordination policy choices may differ from an ex-ante perspective, but in a symmetric Nash-equilibrium the economy inherits the Markov structure of the shock process and is either in the shock state or in steady state. For the case of coordination, Proposition 1 provides the closed-form solution. Figure 1.2 displays the results. The horizontal axis

[23] Eggertsson and Singh (2016) study the accuracy of the loglinear approximate version of the New Keynesian closed-economy model and find the approximate model works well even for a great depression scenario where output falls by some 30 percent. We verify that for our parameter choice the economy's departure from steady state remains within this range, even as no fiscal stabilization takes place.

[24] Replication files are available under http://www.runmycode.org/companion/view/3105.

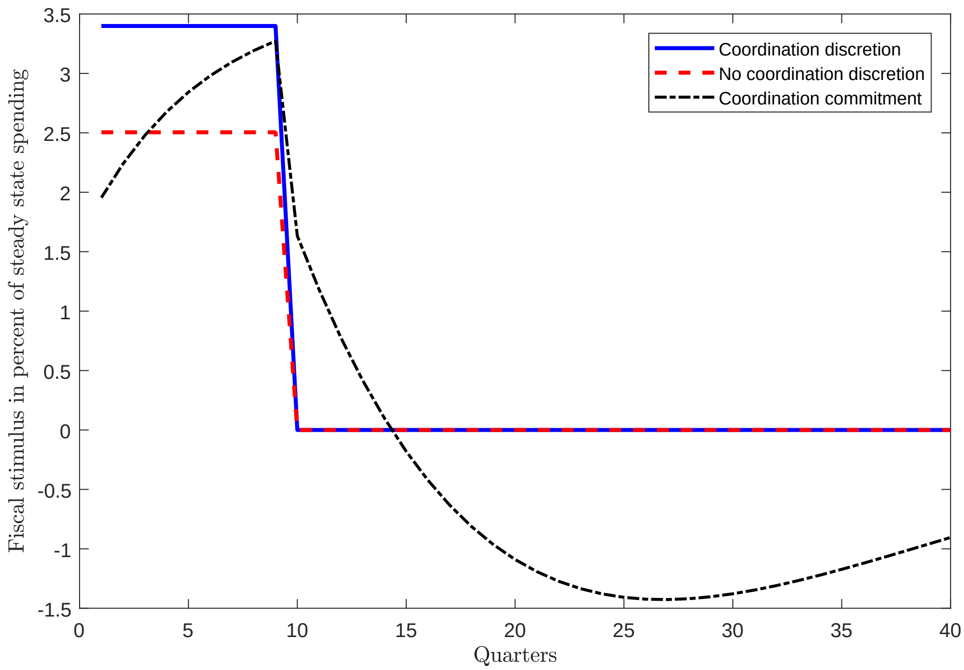

Figure 1.2: Fiscal stimulus at the ELB when ELB ceases to bind in quarter 10, assuming parameter values as given in Table 1.1. Coordination and discretion (solid line), no coordination and discretion (dashed line) and coordination with commitment (dashed-dotted line). Change of spending measured in percent of steady state spending w/ and w/o coordination, respectively.

displays time in quarters. The vertical axis measures the increase of government spending with and without coordination in percentage deviations from the respective steady state. The underlying experiment assumes that while the expected duration of the ELB is 5 quarters, the ELB actually ceases to bind in quarter 10 only. For as long as it binds, the optimal response under coordination is to increase government spending by about 3.5 percent of the steady-state level under coordination (solid line). Instead, in the absence of coordination the optimal response is less aggressive: spending is increased by only 2.5 (dashed line). We conclude that whether there is coordination or not matters quantitatively.

As a benchmark, Figure 1.2 also shows the optimal response in case there is commitment and coordination (dashed-dotted line). To compute the solution for this case we implement the algorithm put forward in Eggertsson and Woodford (2003), see Appendix 1.H. The response under commitment differs from the response under discretion in two ways. First, for as long the ELB binds, government spending is constant under discretion (and above the steady-state level), but increasing under commitment. Such "backloading", as explained in Werning (2012), is an effective way to stabilize the economy at the ELB because of forward-looking price setting: as firms anticipate higher spending and hence higher inflation in the future, they will, all else equal, raise prices already today. This is beneficial in a deflationary environment.[25]

Second, after the ELB ceases to bind, government spending is at the steady-state level under discretion, but reduced below steady state levels under commitment: a "spending reversal". Intuitively, this is beneficial, because after the ELB ceases to bind, a cut of government spending lowers inflation and real interest rates (because monetary policy satisfies the Taylor principle). Under commitment this lowers long-term real interest rates while the ELB binds and thus contributes to stabilizing the economy during the ELB period (Corsetti et al., 2010).

Turning back to the case of discretion, we compute a comprehensive measure of the difference in fiscal stabilization at the ELB: the "stimulus gap", that is, the difference between the optimal spending response without coordination and with coordination, each measured in percentage deviation from the respective steady state ($\hat{g}_L^{*,N} - \hat{g}_L^{*,C}$). Figure 1.3 illustrates that the stimulus gap, measured along the vertical axis, increases strongly (in absolute value) in the probability that the ELB binds for another period μ, measured along the horizontal axis. Whether fiscal policies are coordinated or not hardly matters if the expected duration of the ELB episode is short. However, for larger values of μ the stimulus gap is sizable and reaches up to (minus) 4 percentage points.

1.4.2 Sensitivity analysis

We now explore how the stimulus gap varies with other important model parameters. The panels in Figure 1.4 vary one parameter at a time when computing the stimulus gap. We first consider alternative values for the discount factor β. It determines how the dynamics of the terms of trade (as endogenous state variable) impact optimal policy in

[25] Werning (2012) studies a closed-economy model and so does Schmidt (2017) who finds that an activist fiscal policymaker who cares less about the stabilization of public consumption than society may partly correct for discretionary authorities' inability to exploit the "expectations channel".

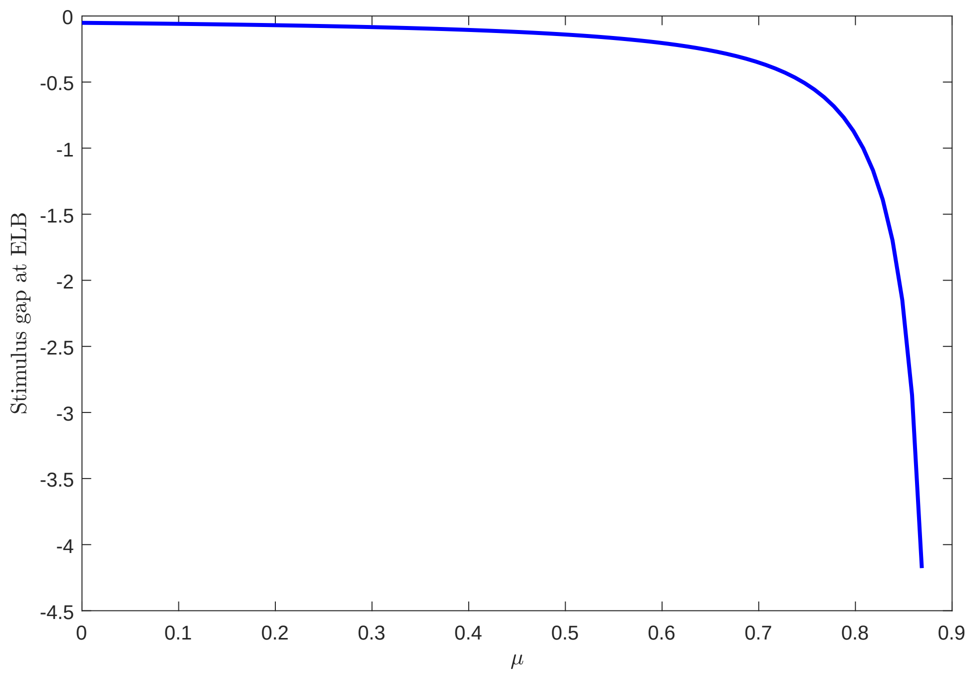

Figure 1.3: Fiscal stimulus gap at the ELB. Difference of optimal increase of government consumption without and with coordination, measured in percentage points along the vertical axis ($\hat{g}_L^{*,N} - \hat{g}_L^{*,C}$), horizontal axis measures probability that the ELB remains binding for another period (μ). We only consider parameter values for which the equilibrium is locally unique.

the absence of coordination (see the discussion in Section 1.3.2 above).[26] The upper-left panel of Figure 1.4 shows the stimulus gap as a function of β. We find that the stimulus gap increases (in absolute value) as β increases, because in the absence of coordination the desire to avoid a persistent appreciation of the terms of trade increases with β.

The upper-right panel of Figure 1.4 shows the stimulus gap as a function of the degree of price stickiness θ. We find that the lower the degree of price stickiness, the larger the difference between the optimal policy under coordination and without. To understand

[26] At the same time β matters also for the slope of the Phillips curve. Yet, as we vary the value of β, we keep κ constant (just like all other parameters) in order to focus on the role of the dynamic terms of trade effect.

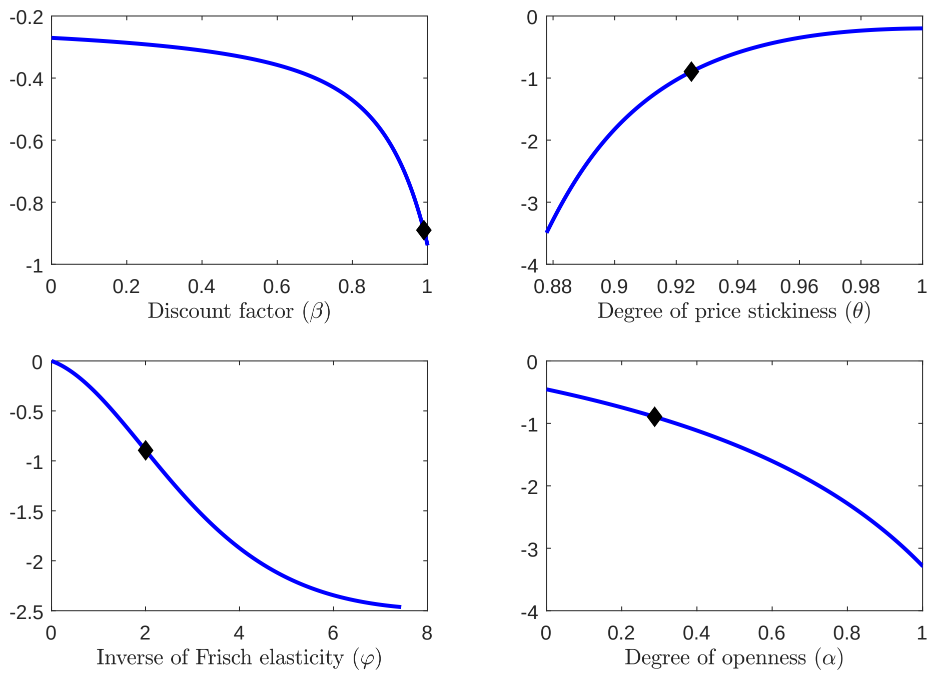

Figure 1.4: Fiscal stimulus gap at the ELB, measured in percentage points along the vertical axes ($\hat{g}_L^{*,N} - \hat{g}_L^{*,C}$), as a function of model parameters, measured along the horizontal axis. We only consider parameter values for which the equilibrium is locally unique. A diamond indicates parameter values for the baseline.

this finding, note that inflation responds more strongly to higher government spending if prices are more flexible. Higher inflation at the union level reduces real interest rates and thus stimulates aggregate demand at the union level. At the country level, instead, higher inflation appreciates the terms of trade and thus reduces the demand for domestically produced goods. Hence, the more flexible prices, the more negative the stimulus gap.

The lower-left panel of Figure 1.4 shows the stimulus gap conditional on the inverse of the Frisch elasticity of labor supply φ. As the Frisch elasticity declines (that is, as φ increases), marginal costs and, hence, inflation respond more strongly to higher government spending. Put differently, the Phillips curve becomes steeper as φ increases, just like when θ declines, see equation (1.10). We therefore find the stimulus gap more negative, the larger the Frisch elasticity.

The lower-right panel of Figure 1.4 shows the stimulus gap conditional on openness parameter α. It is possible to show that, the higher the degree of openness, the stronger the impact of government spending on the terms of trade. Consequently, in a more open economy policymakers seek to avoid a stronger appreciation of the terms of trade in the midst of a severe recession and provide a lower stimulus in the absence of policy coordination. Note that the stimulus gap does not vanish in the closed economy limit ($\alpha = 0$) since monetary regimes still differ. While there is an implicit price-level target in place at the country-level, the price level features a unit root at the union level. This difference has a strong bearing on the transmission of fiscal policy (Corsetti et al., 2013a).

1.4.3 Distortionary taxes

Our baseline scenario assumes that taxes are lump-sum. We now investigate to what extent our results hinge on this assumption. For this purpose, we continue to assume that the government runs a balanced budget, but in order to do so, it adjusts labor income taxes which are denoted by $\tau_{N,t}^i$. We replace the government budget constraint (1.2) with the following equation:

$$T^i + \tau_{N,t}^i W_t^i N_t^i = P_t^i G_t^i + \tau^i W_t^i N_t^i.$$

The lump-sum tax T^i is still used to finance the employment subsidy ($\tau^i W_t^i N_t^i$) in order to decentralize the social planner steady states. Linearizing the government budget constraint gives the following equation after substituting for the real wage and labor:

$$\left(\frac{1}{1-\gamma} + (1+\varphi)\right)\hat{y}_t^i + \frac{1}{1-\tau_N}\hat{\tau}_{N,t}^i = \frac{1}{1-\gamma}\hat{g}_t^i,$$

where τ_N denotes the steady state labor income tax rate. Further, with distortionary taxes the New Keynesian Phillips curve at the country level changes to:

$$\pi_t^i = \beta E_t\{\pi_{t+1}^i\} + \lambda\left(\frac{1}{1-\gamma} + \varphi\right)\hat{y}_t^i - \frac{\lambda\gamma}{1-\gamma}\hat{g}_t^i + \frac{\lambda\tau_N}{1-\tau_N}\hat{\tau}_{N,t}^i. \tag{1.25}$$

A corresponding equation also holds at the union level.[27]

[27]The modifications of the rest of the model are moderate. A formal exposition is available upon request.

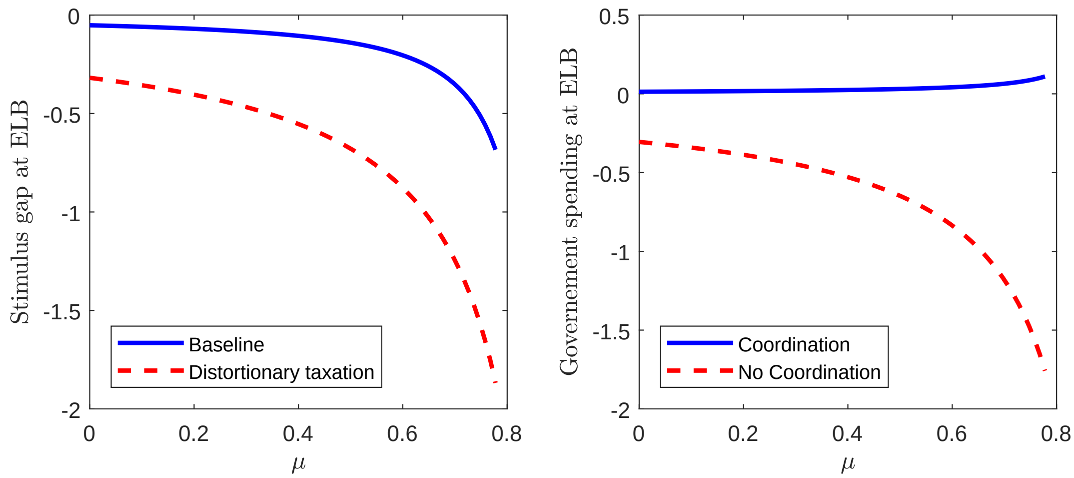

Figure 1.5: Left panel shows stimulus gap at the ELB for lump sum taxes as in baseline (solid line) and for distortionary taxes (dashed line). Right panel shows government spending with coordination (solid line) and without (dashed line), measured in percent of respective steady state. Horizontal axes measure the probability that the ELB remains binding for another period (μ). We only consider parameter values for which the equilibrium is locally unique.

Figure 1.5 shows the result: the left panel contrasts the stimulus gap for distortionary taxes (dashed line) with the stimulus gap in the baseline case when taxes are lump-sum (solid line). As before, the horizontal axis measures the probability that the ELB remains binding for another period. It turns out that the stimulus gap is considerably larger (in absolute value) when taxes are distortionary. To understand this finding, note that, all else equal, raising taxes is inflationary, see equation (1.25). Higher taxes at the union-level can therefore be expansionary at the ELB: because (expected) inflation is not met by a higher interest rate, the real interest rate declines and the fiscal multiplier increases at the ELB, as emphasized by Eggertsson (2011). At the country level, however, the inflationary effect of higher distortionary taxes appreciates the terms of trade. Hence, the difference between the optimal stimulus under coordination and without increases if taxes are distortionary.

The right panel of Figure 1.5 illustrates further the economic mechanism at play. It displays the optimal response of government spending to the crisis, both for coordination and without, in percent of the respective steady state.[28] Under coordination (solid line), government spending is raised only mildly because of the inflationary effects of increasing the income tax. In the absence of coordination (dashed line), however, government spending is *cut* because this brings about a tax cut and thus weakens the terms of trade if taxes are distortionary.

1.4.4 The level of government spending and welfare

So far we have focused on the percentage change of government spending at the ELB. The stimulus gap, in particular, is computed as the percentage point difference of the stimulus with and without coordination. As such it provides a measure for the difference in the degree of fiscal stabilization and it turns out that the gap can be sizeable: it easily amounts to 4 percentage points, as shown in Figure 1.3. Government spending, in other words, is raised more aggressively under coordination. However, we have established in Section 1.3.1 above that in steady state the *level* of government spending is lower under coordination than without.

In what follows we therefore compare the optimal *level* of government spending at the ELB. The left panel of Figure 1.6 shows the result. The vertical axis measures the optimal level of government consumption with (solid line) and without (dashed line) coordination. Along the horizontal axis we measure again the expected duration of the ELB episode. As before, results are based on the parameter values listed in Table 1.1 above. We find that the steady-state effect dominates: the level of spending without coordination exceeds the optimal level with coordination for all values of μ for which there is a unique equilibrium. Nevertheless, the distance between the optimal level of government expenditure with and without coordination becomes smaller as the expected duration of the ELB episode increases.[29]

Finally, we assess the welfare costs of a lack of coordination at the ELB. Specifically, we compute the compensation in terms of consumption which is required to make the household indifferent between the optimal stimulus under coordination and the stimulus which is optimal in the absence of coordination. For this purpose we assume that the

[28] Compared to Figure 1.3 we plot the stimulus gap in Figure 1.5 for a smaller range of μ because the inflationary effects of an increased income tax imply positive interest rates for high values of μ.

[29] For alternative parameter values the optimal level of spending without coordination actually falls short of the optimal level with coordination. For instance, assuming $\varphi = 0.2$ and $\theta = 0.5$ we find this to be the case, provided the ELB episode is expected to be sufficiently long lasting.

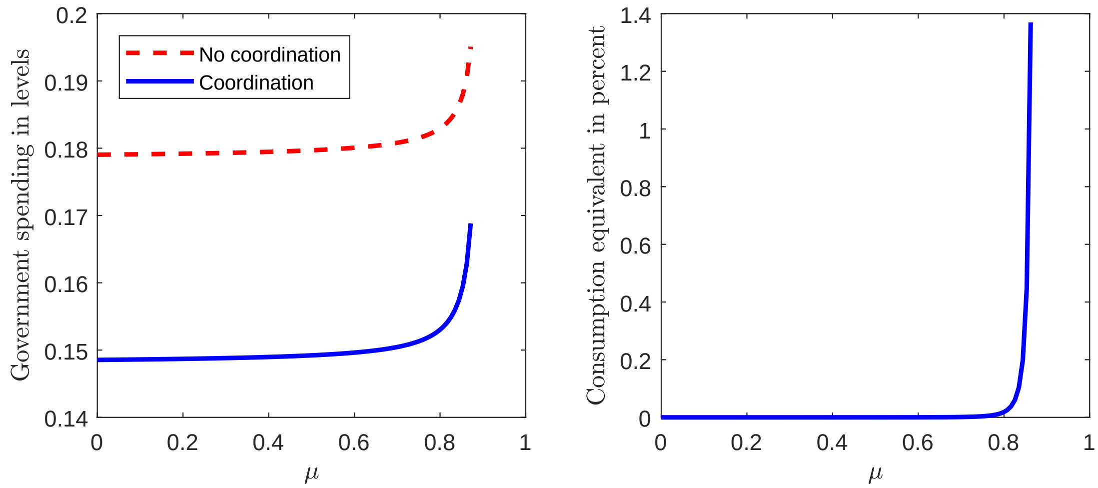

Figure 1.6: Left panel shows optimal level of government spending under coordination (solid line) and without (dashed line): G_t^C and G_t^N. Right panel shows consumption equivalent which compensates for the lack of coordination for as long as the ELB binds (percent of steady state consumption with coordination, see Appendix 1.I). Horizontal axes measure the probability that the ELB remains binding for another period (μ). We only consider parameter values for which the equilibrium is locally unique.

economy is initially at the same steady state, namely the steady state which obtains under coordination. We then compute the consumption equivalent by comparing the percentage increase of government spending which is optimal under coordination with the percentage increase which would have been optimal absent coordination. Importantly, we assume that the household receives the compensation only for as long as the ELB binds (see Appendix 1.I for details).[30]

The right panel of Figure 1.6 displays the result. The vertical axis measures the consumption equivalent in percent of steady-state consumption. The horizontal axis measures the probability μ that the ELB remains binding for another period. As shown above, the stimulus gap is negligible for low values of μ. For this reason the consumption equivalent is also negligible in this parameter range. However, if the ELB is expected to bind for a relatively long period, the benefits of coordination increase substantially. For example, when the expected duration is 7 quarters ($\mu = 0.8571$) the consumption equivalent

[30] Benigno and Benigno (2006) stress the importance of comparing optimal (monetary) policy with and without coordination around the same steady state.

amounts to about 0.69%.[31] However, we stress once more that the consumption equivalent is received only for as long the ELB binds, not permanently.

1.5 Conclusion

In the context of the global financial crisis fiscal policy has been rediscovered as a stabilization tool. Central in this context is that monetary policy has become constrained by the ELB on nominal interest rates. It not only seems natural to turn to fiscal policy for additional support, it has also been established that fiscal policy is likely to be particularly effective under these circumstances. Against this background, we consider a currency union where a common monetary policy operates jointly with many fiscal policies. Assuming that the common monetary policy is unable to stabilize area-wide inflation and output because of the ELB constraint, we ask whether there is a need to coordinate government spending policies across the member states of the union.

Our analysis is based on the model of Galí and Monacelli (2008) which we extend in order to account for the ELB and the absence of fiscal policy coordination. Absent coordination, we find that policymakers use fiscal policy less aggressively than under coordination. The resulting fiscal "stimulus gap" can be sizeable and the welfare gains from coordination appear non trivial. Intuitively, what makes local policymakers reluctant to use fiscal policy aggressively, is that unilateral fiscal stimulus appreciates the terms of trade and is therefore less effective in stabilizing the local economy. However, if enacted across the entire union, fiscal stimulus leaves the terms of trade of any single country unaffected and the union-wide inflationary impulse stimulates economic activity at the ELB.

This result is specific to currency unions at the ELB. If the ELB does not bind, monetary policy is perfectly able to stabilize union-wide output and inflation. Hence, in this sense there is no need to coordinate fiscal policies. If the ELB binds, but exchange rates are flexible, the terms of trade depreciate in response to a unilateral fiscal expansion. Hence, in this case, the mechanism which drives our result changes fundamentally. And indeed, Cook and Devereux (2011) consider a two-country model with flexible exchange rates and find

[31] In the right panel of Figure 1.6 we restrict the range of μ such that the equilibrium drop of output and inflation remain empirically plausible. At the maximum value of μ considered in Figure 1.6, the annualized drop of output and inflation is 16% and 2.6% in the absence of coordination. At this point, the government spending multiplier at the union level amounts to 3.4. The multiplier and the welfare gains from coordination increase strongly in μ; since we consider such a scenario as implausible we restrict the range of μ accordingly.

that the case for fiscal coordination is weak, even as the ELB binds. For these reasons, the case of a currency union at the ELB is a special one. That said, it is certainly a relevant one: our results are consistent with the view that the euro area did not receive sufficient fiscal stimulus during the recent crisis. Still, a more comprehensive assessment would also need to account for issues related to sovereign risk, an aspect from which we abstract in our analysis. We leave this for future research.

1.A Steady state in the absence of coordination

1.A.1 Planner problem

The risk sharing condition and the market-clearing condition in (1.19) imply the following equation:

$$C^i = (Y^i - G^i)^{1-\alpha}(C^*)^\alpha. \tag{1.A.1}$$

The Lagrangian associated with problem (1.19) is thus given by:

$$\mathcal{L} = (1-\chi)\log C^i + \chi \log G^i - \frac{\left(N^i\right)^{1+\varphi}}{1+\varphi} + \Lambda(C^i - (N^i - G^i)^{1-\alpha}(C^*)^\alpha).$$

First order conditions are given by:

$$\frac{\partial \mathcal{L}}{\partial C^i} = (1-\chi)\frac{1}{C^i} + \Lambda = 0 \tag{1.A.2}$$

$$\begin{aligned} \frac{\partial \mathcal{L}}{\partial N^i} &= -(N^i)^\varphi - \Lambda(1-\alpha)(N^i - G^i)^{-\alpha}(C^*)^\alpha \\ &\overset{(1.A.1)}{=} -(N^i)^\varphi - \Lambda(1-\alpha)\frac{C^i}{Y^i - G^i} = 0 \end{aligned} \tag{1.A.3}$$

$$\begin{aligned} \frac{\partial \mathcal{L}}{\partial G^i} &= \chi\frac{1}{G^i} + \Lambda(1-\alpha)(N^i - G^i)^{-\alpha}(C^*)^\alpha \\ &= \chi\frac{1}{G^i} + \Lambda(1-\alpha)\frac{C^i}{Y^i - G^i} = 0. \end{aligned} \tag{1.A.4}$$

Combine (1.A.3) and (1.A.4) to get:

$$(N^i)^\varphi = \chi\frac{1}{G^i}. \tag{1.A.5}$$

Further, combine (1.A.2) and (1.A.4):

$$\frac{\chi}{1-\chi}\frac{C^i}{G^i} = (1-\alpha)\frac{C^i}{Y^i - G^i}.$$

Which can be rearranged to:

$$G^i = \frac{\chi}{1-\chi}\left((1-\alpha) + \frac{\chi}{1-\chi}\right)^{-1} Y^i. \tag{1.A.6}$$

It thus follows for the absence of coordination that in each country $i \in [0,1]$ we have

$$\left(\frac{G}{Y}\right)^N = \frac{\chi}{(1-\alpha)(1-\chi)+\chi}. \tag{1.A.7}$$

Since in a symmetric steady state $Y^i = N^i$, combining (1.A.5) and (1.A.7) gives for each country:

$$N^N = Y^N = [(1-\alpha)(1-\chi)+\chi]^{\frac{1}{1+\varphi}}. \tag{1.A.8}$$

Further, it holds that $G^C < G^N$ since

$$G^C = \chi < \frac{\chi}{(1-\alpha)(1-\chi)+\chi}[(1-\alpha)(1-\chi)+\chi]^{\frac{1}{1+\varphi}} = G^N.$$

1.A.2 Decentralization of the planner solution in steady state

In the following we show how the planner allocation in the absence of coordination can be decentralized in a symmetric zero-inflation steady state. Unless offset by the employment subsidy, firms in country i choose a constant mark-up over marginal costs MC^i, which can be expressed as (see equation (41) in Galí and Monacelli, 2008):

$$1 - \frac{1}{\varepsilon} = MC^i = \frac{1-\tau^i}{1-\chi}(N^i)^{1+\varphi}\left(1 - \frac{G^i}{Y^i}\right).$$

In order to decentralize the planner solution government consumption, G^i, has to be set according to (1.A.6). Solving the resulting expression for N^i gives:

$$(N^i)^{1+\varphi} = \left(1 - \frac{1}{\varepsilon}\right) \frac{1-\chi}{1-\tau^i} \left(1 - \frac{\chi}{1-\chi} \left((1-\alpha) + \frac{\chi}{1-\chi}\right)^{-1}\right)^{-1}.$$

Further, in the absence of coordination the following subsidy has to be chosen in each country i

$$1 - \tau^N = \left(1 - \frac{1}{\varepsilon}\right)(1-\alpha)^{-1}, \tag{1.A.9}$$

such that the social planner solution without coordination is decentralized in a symmetric zero-inflation steady state.

1.B Deriving the welfare function without coordination

In the absence of coordination there are linear terms in a second-order approximation to household utility. We follow the approach of Benigno and Woodford (2006) and substitute for the linear terms using a second order approximation to the market-clearing condition. In the following we drop the country index i for simplicity and approximate the percentage deviation of a generic variable X_t from its steady state X by

$$\frac{X_t - X}{X} \approx \hat{x}_t + \frac{1}{2}\hat{x}_t^2,$$

where $\hat{x}_t = x_t - x$ and $x_t = \log X_t$.

1.B.1 Second order approximation to the goods market clearing condition

The market-clearing condition is given by $Y_t^i = C_t^i (S_t^i)^\alpha + G_t^i$. Taking logs and rearranging gives:

$$\log C_t^i = \log(Y_t^i - G_t^i) - \alpha \log S_t^i.$$

A second-order approximation to the above equation gives:

$$\hat{c}_t^i \approx \frac{1}{1-\gamma}\left(\hat{y}_t^i - \gamma \hat{g}_t^i\right) - \frac{1}{2}\frac{\gamma}{(1-\gamma)^2}(\hat{g}_t^i - \hat{y}_t^i)^2 - \alpha s_t^i. \tag{1.B.1}$$

Combining the above equation with the risk-sharing condition (1.4) yields:

$$0 \approx \frac{1}{1-\gamma}\left(\hat{y}_t^i - \gamma \hat{g}_t^i\right) - \frac{1}{2}\frac{\gamma}{(1-\gamma)^2}(\hat{g}_t^i - \hat{y}_t^i)^2 - s_t^i + t.i.p. \tag{1.B.2}$$

where $(t.i.p.)$ captures terms independent of policy, namely $\hat{c}_t^*$ since it evolves exogenously for a given member of the currency union. For future reference, we define:

$$A_y \equiv \frac{1}{1-\gamma}; \quad A_g \equiv -\frac{\gamma}{1-\gamma}; \quad A_s \equiv -1. \tag{1.B.3}$$

1.B.2 Second order approximation to utility

Utility in country i, $U_t^i = U(C_t^i, G_t^i, N_t^i)$, is additively separable in its arguments. A second order approximation around a generic steady state C^i, G^i, N^i therefore gives:

$$\begin{aligned} U_t^i - U^i \approx\ & U_C^i C^i \left(\frac{C_t^i - C^i}{C^i}\right) + U_G^i G^i \left(\frac{G_t^i - G^i}{G^i}\right) + U_N^i N^i \left(\frac{N_t^i - N^i}{N^i}\right) \\ & + \frac{1}{2}U_{CC}^i (C^i)^2 \left(\frac{C_t^i - C^i}{C^i}\right)^2 + \frac{1}{2}U_{GG}^i (G^i)^2 \left(\frac{G_t^i - G^i}{G^i}\right)^2 \\ & + \frac{1}{2}U_{NN}^i (N^i)^2 \left(\frac{N_t^i - N^i}{N^i}\right)^2. \end{aligned}$$

Rewriting the expression in terms of log deviations the above approximation becomes:

$$\begin{aligned} U_t^i - U^i \approx\ & U_C^i C^i \left(\hat{c}_t^i + \frac{1}{2}(\hat{c}_t^i)^2\right) + U_G^i G^i \left(\hat{g}_t^i + \frac{1}{2}(\hat{g}_t^i)^2\right) + U_N^i N^i \left(\hat{n}_t^i + \frac{1}{2}(\hat{n}_t^i)^2\right) \\ & + \frac{1}{2}U_{CC}^i (C^i)^2 (\hat{c}_t^i)^2 + \frac{1}{2}U_{GG}^i (G^i)^2 (\hat{g}_t^i)^2 + \frac{1}{2}U_{NN}^i (N^i)^2 (\hat{n}_t^i)^2. \end{aligned}$$

Rearranging:

$$\begin{aligned} U_t^i - U^i \approx\ & U_C^i C^i \left(\hat{c}_t^i + \frac{1}{2}\left(1 + \frac{U_{CC}^i C^i}{U_C^i}\right)(\hat{c}_t^i)^2\right) + U_G^i G^i \left(\hat{g}_t^i + \frac{1}{2}\left(1 + \frac{U_{GG}^i G^i}{U_G^i}\right)(\hat{g}_t^i)^2\right) \\ & + U_N^i N^i \left(\hat{n}_t^i + \frac{1}{2}\left(1 + \frac{U_{NN}^i N^i}{U_N^i}\right)(\hat{n}_t^i)^2\right). \end{aligned}$$

Defining further: $\sigma \equiv -\frac{U_{CC}C}{U_C}$, $\sigma_g \equiv -\frac{U_{GG}G}{U_G}$ and $\sigma_n \equiv \frac{U_{NN}N}{U_N}$ yields

$$U_t^i - U^i \approx \; U_C^i C^i \left(\hat{c}_t^i + \frac{1}{2}\left(1-\sigma\right)\left(\hat{c}_t^i\right)^2 \right) + U_G^i G^i \left(\hat{g}_t^i + \frac{1}{2}\left(1-\sigma_g\right)\left(\hat{g}_t^i\right)^2 \right) + U_N^i N^i \left(\hat{n}_t^i + \frac{1}{2}\left(1+\sigma_n\right)\left(\hat{n}_t^i\right)^2 \right).$$

Since utility is given by

$$U_t^i = (1-\chi)\log C_t^i + \chi \log G_t^i - \frac{(N_t^i)^{1+\varphi}}{1+\varphi},$$

the above defined parameters become: $\sigma = \sigma_g = 1$ while $\sigma_n = \varphi$, such that we get:

$$\frac{U_t^i - U^i}{U_C^i C^i} \approx \; \hat{c}_t^i + \frac{U_G^i G^i}{U_C^i C^i}\hat{g}_t^i + \frac{U_N^i N^i}{U_C^i C^i}\left(\hat{n}_t^i + \frac{1}{2}(1+\varphi)(\hat{n}_t^i)^2 \right).$$

Because of monopolistic competition firms charge a markup over marginal costs. If not offset by a certain value for the labor subsidy there will be a wedge Φ between the marginal rate of substitution and the marginal product of labor (MPN) in steady state (see, for instance, Galí, 2008, p.106):

$$-\frac{U_N^i}{U_C^i} = \mathrm{MPN}^i(1-\Phi). \tag{1.B.4}$$

In our setup we have $\mathrm{MPN} = Y/N$. Therefore

$$\frac{U_N^i}{U_C^i}\frac{N^i}{C^i} = -\frac{1}{1-\gamma}(1-\Phi),$$

with $1-\gamma = C/Y$. Making use of the above expression and the one for $\frac{U_G G}{U_C C}$ under the assumed utility function, we can rewrite the approximation to utility as:

$$\frac{U_t^i - U^i}{U_C^i C^i} \approx \; \hat{c}_t^i + \frac{\chi}{1-\chi}\hat{g}_t^i - \frac{1-\Phi}{1-\gamma}\left(\hat{n}_t^i + \frac{1}{2}(1+\varphi)(\hat{n}_t^i)^2 \right).$$

Further, we use equation (1.B.1) in order to substitute for $\hat{c}_t^i$. Therefore

$$\begin{aligned}\frac{U_t^i - U^i}{U_C^i C^i} \approx\ & \frac{1}{1-\gamma}\left(\hat{y}_t^i - \gamma \hat{g}_t^i\right) - \frac{1}{2}\frac{\gamma}{(1-\gamma)^2}(\hat{g}_t^i - \hat{y}_t^i)^2 - \alpha s_t^i \\ & + \frac{\chi}{1-\chi}\hat{g}_t^i - \frac{1-\Phi}{1-\gamma}\left(\hat{n}_t^i + \frac{1}{2}(1+\varphi)(\hat{n}_t^i)^2\right) + t.i.p.\end{aligned}$$

In order to substitute for $\hat{n}_t^i$, it can be shown that aggregate labor demand is given by $N_t^i = Y_t^i \int_0^1 \left(\frac{P_t^i(j)}{P_t^i}\right)^{-\varepsilon} dj$, see Galí and Monacelli (2008). Define $z_t^i \equiv \log \int_0^1 \left(\frac{P_t^i(j)}{P_t^i}\right)^{-\varepsilon} dj$. Thus, it holds around a symmetric steady state that:

$$\hat{n}_t^i = \hat{y}_t^i + z_t^i.$$

Further it can be shown that z_t^i is of second order with $z_t^i \approx \frac{1}{2}\frac{\varepsilon}{\lambda}(\pi_t^i)^2$, see again Galí and Monacelli (2008). Finally, the approximation to utility can be expressed as:

$$\begin{aligned}\frac{U_t^i - U^i}{U_C^i C^i} \approx\ & \frac{\Phi}{1-\gamma}\hat{y}_t^i + \left(\frac{\chi}{1-\chi} - \frac{\gamma}{1-\gamma}\right)\hat{g}_t^i - \alpha s_t^i \\ & - \frac{1}{2}\frac{\gamma}{(1-\gamma)^2}(\hat{g}_t^i - \hat{y}_t^i)^2 - \frac{1}{2}\frac{1-\Phi}{1-\gamma}(1+\varphi)(\hat{y}_t^i)^2 - \frac{1}{2}\frac{1-\Phi}{1-\gamma}\frac{\varepsilon}{\lambda}(\pi_t^i)^2 + t.i.p.\end{aligned} \tag{1.B.5}$$

For future reference we define:

$$B_y \equiv \frac{\Phi}{1-\gamma}; \quad B_g \equiv \frac{\chi}{1-\chi} - \frac{\gamma}{1-\gamma}; \quad B_s \equiv -\alpha. \tag{1.B.6}$$

1.B.3 The welfare function—substituting for the linear terms

Absent coordination, government spending and the employment subsidy, τ, are not chosen efficiently such that $\gamma = \gamma^N \neq \chi$ and $\Phi \neq 0$. Specifically, in a symmetric steady state the distortion Φ is given by (see Galí, 2008, p.73 and p.106):

$$\Phi = 1 - \frac{1}{\frac{\varepsilon}{\varepsilon-1}(1-\tau)}.$$

Inserting for the subsidy according to (1.A.9) yields:

$$\Phi = \alpha.$$

By inserting for γ^N and Φ in (1.B.3) and (1.B.6) we get:

$$A_y = \frac{(1-\chi)(1-\alpha)+\chi}{(1-\chi)(1-\alpha)}; \quad A_g = -\frac{\chi}{(1-\chi)(1-\alpha)}; \quad A_s = -1;$$
$$B_y = \frac{\alpha[(1-\chi)(1-\alpha)+\chi]}{(1-\chi)(1-\alpha)}; \quad B_g = -\frac{\alpha\chi}{(1-\chi)(1-\alpha)}; \quad B_s = -\alpha.$$

Thus, it is easily seen that subtracting α times condition (1.B.2) from (1.B.5)—both evaluated at the steady state in the absence of coordination—removes the linear terms from the approximation to utility. As a result, the welfare function is given by:

$$\mathbb{W}_t^N \approx -\frac{1}{2}\left(\frac{\varepsilon}{\lambda}(\pi_t^i)^2 + (1+\varphi)(\hat{y}_t^i)^2 + \frac{\gamma^N}{1-\gamma^N}\left(\hat{g}_t^i - \hat{y}_t^i\right)^2\right) + t.i.p.$$

1.C Optimal policy with coordination

The Lagrangian associated with problem (1.21) is given by

$$\begin{aligned}\mathcal{L}_t = &-\frac{1}{2}\left(\frac{\varepsilon}{\lambda}(\pi_t^*)^2 + (1+\varphi)(\hat{y}_t^*)^2 + \frac{\gamma^C}{1-\gamma^C}(\hat{g}_t^* - \hat{y}_t^*)^2\right) \\ &+ \xi_{0,t}^*\left[\pi_t^* - \lambda\left(\frac{1}{1-\gamma^C} + \varphi\right)\hat{y}_t^* + \frac{\lambda\gamma^C}{1-\gamma^C}\hat{g}_t^* - \nu_{0,t}^*\right] \\ &+ \xi_{1,t}^*\left[-(\hat{y}_t^* - \gamma^C\hat{g}_t^*) + \nu_{1,t}^*\right].\end{aligned}$$

The Kuhn-Tucker conditions read as follows:

$$\frac{\partial\mathcal{L}_t}{\partial\pi_t^*} = -\frac{\varepsilon}{\lambda}\pi_t^{*,C} + \xi_{0,t}^* = 0 \tag{1.C.1}$$

$$\frac{\partial\mathcal{L}_t}{\partial\hat{y}_t^*} = -(1+\varphi)\hat{y}_t^{*,C} + \frac{\gamma^C}{1-\gamma^C}(\hat{g}_t^{*,C} - \hat{y}_t^{*,C}) - \lambda\left(\frac{1}{1-\gamma^C} + \varphi\right)\xi_{0,t}^* - \xi_{1,t}^* = 0 \tag{1.C.2}$$

$$\frac{\partial\mathcal{L}_t}{\partial\hat{g}_t^*} = -\frac{\gamma^C}{1-\gamma^C}(\hat{g}_t^{*,C} - \hat{y}_t^{*,C}) + \frac{\lambda\gamma^C}{1-\gamma^C}\xi_{0,t}^* + \gamma^C\xi_{1,t}^* = 0 \tag{1.C.3}$$

$$\xi_{1,t}^*(-(\hat{y}_t^{*,C} - \gamma^C\hat{g}_t^{*,C}) + \nu_{1,t}^*) = 0$$

$$\xi_{1,t}^* \geq 0; \;\; -(\hat{y}_t^{*,C} - \gamma^C\hat{g}_t^{*,C}) + \nu_{1,t}^* \geq 0.$$

Case 1 The effective-lower-bound constraint does not bind: $\xi^*_{1,t} = 0$ and $-(\hat{y}^{*,C}_t - \gamma^C \hat{g}^{*,C}_t) + \nu^*_{1,t} \geq 0$. Equation (1.C.3) thus implies for $\xi^*_{0,t}$

$$\xi^*_{0,t} = \frac{1}{\lambda}(\hat{g}^{*,C}_t - \hat{y}^{*,C}_t). \tag{1.C.4}$$

Inserting for $\xi^*_{0,t}$ in (1.C.2) yields

$$\hat{g}^{*,C}_t = 0.$$

Using this in (1.C.4) and combining it with (1.C.1) gives

$$\hat{y}^{*,C}_t = -\varepsilon \pi^{*,C}_t.$$

Therefore, when it is optimal to stabilize output at steady state inflation should be zero and vice versa. And indeed, considering the functional form of the welfare function it is clear that $\pi^{*,C}_t = \hat{y}^{*,C}_t = \hat{g}^{*,C}_t = 0$ is the global maximum.

Case 2 The effective-lower-bound constraint binds: $\xi^*_{1,t} > 0$ and $-(\hat{y}^{*,C}_t - \gamma^C \hat{g}^{*,C}_t) + \nu^*_{1,t} = 0$. Rearrange (1.C.3):

$$-\xi^*_{1,t} = -\frac{1}{1-\gamma^C}(\hat{g}^{*,C}_t - \hat{y}^{*,C}_t) + \frac{\lambda}{1-\gamma^C}\xi^*_{0,t}.$$

Combining it with (1.C.2) yields after rearranging

$$\xi^*_{0,t} = -\frac{1}{\lambda}\hat{y}^{*,C}_t - \frac{1}{\lambda\varphi}\hat{g}^{*,C}_t.$$

Inserting for $\xi^*_{0,t}$ in (1.C.1) gives

$$\pi^{*,C}_t + \frac{1}{\varepsilon}\hat{y}^{*,C}_t = -\psi^C_g \hat{g}^{*,C}_t, \tag{1.C.5}$$

where $\psi^C_g \equiv \frac{1}{\varepsilon\varphi}$.

1.D Proof of proposition 1

In proposition 1 we state the solution for optimal government consumption at the ELB with coordination. Under the ELB scenario under consideration, we get a two-state solution (see Section 1.2.2). Optimal policy at the ELB is determined by equations (1.14), (1.16) and (1.C.5) with $\psi_t = 0$. Further applying $\frac{\kappa\bar{\sigma}\gamma}{\bar{\sigma}+\varphi} = \frac{\lambda\gamma}{1-\gamma}$ in (1.14) and (1.16), the three equations read:

$$\pi_L^{*,C} = \frac{1}{(1-\beta\mu)}\kappa^C \hat{y}_L^{*,C} - \frac{1}{(1-\beta\mu)}\frac{\lambda\gamma^C}{(1-\gamma^C)}\hat{g}_L^{*,C} \tag{1.D.1}$$

$$\hat{y}_L^{*,C} = \frac{(1-\gamma^C)(1-\beta\mu)}{(1-\mu)(1-\beta\mu)-(1-\gamma^C)\mu\kappa^C} r_L + \frac{(1-\mu)(1-\beta\mu)\gamma^C - \mu\lambda\gamma^C}{(1-\mu)(1-\beta\mu)-(1-\gamma^C)\mu\kappa^C}\hat{g}_L^{*,C} \tag{1.D.2}$$

$$-\psi_g^C \hat{g}_L^{*,C} = \pi_L^{*,C} + \frac{1}{\varepsilon}\hat{y}_L^{*,C} \tag{1.D.3}$$

Equations (1.D.1) and (1.D.3) imply:

$$\hat{y}_L^{*,C} = \frac{\frac{\varepsilon\lambda\gamma^C}{1-\gamma^C} - (1-\beta\mu)\psi_g^C\varepsilon}{\kappa^C\varepsilon + (1-\beta\mu)}\hat{g}_L^{*,C}. \tag{1.D.4}$$

Combining (1.D.2) with (1.D.4) and solving for $\hat{g}_L^{*,C}$, we get:

$$\hat{g}_L^{*,C} = -\Theta^C r_L,$$

with

$$\Theta^C \equiv \frac{(1-\gamma^C)\left(\kappa^C\varepsilon + (1-\beta\mu)\right)}{\varepsilon\psi_g^C\Gamma_1^C + (1-\mu)\gamma^C\lambda\varphi\varepsilon + \gamma^C\Gamma_2^C},$$

where $\Gamma_1^C \equiv (1-\mu)(1-\beta\mu) - (1-\gamma^C)\mu\kappa^C$ and $\Gamma_2^C \equiv (1-\mu)(1-\beta\mu) - \mu\lambda$. Since we consider only uniquely determined equilibria (see the ELB scenario in Section 1.2.2), Γ_1^C and Γ_2^C are positive. All other expressions in Θ^C are non-negative. Hence, $\Theta^C > 0$. ■

1.E Optimal policy in the absence of coordination

The Lagrangian associated with problem (1.23) is given by

$$
\begin{aligned}
\mathcal{L}_t = & -\frac{1}{2}\left(\frac{\varepsilon}{\lambda}(\pi_t^i)^2 + (1+\varphi)(\hat{y}_t^i)^2 + \frac{\gamma^N}{1-\gamma^N}\left(\hat{g}_t^i - \hat{y}_t^i\right)^2\right) + \beta E_t V(s_t^i, \pi_{t+1}^*, \hat{y}_{t+1}^*, \hat{g}_{t+1}^*) \\
& + m_{1,t}\left[\hat{y}_t^i - \hat{y}_t^* - \gamma^N \hat{g}_t^i + \gamma^N \hat{g}_t^* - (1-\gamma^N)s_t^i\right] \\
& + m_{2,t}\left[\pi_t^i - \beta E_t\{\pi_{t+1}^i\} - \lambda\left(\frac{1}{1-\gamma^N} + \varphi\right)\hat{y}_t^i + \lambda\frac{\gamma^N}{1-\gamma^N}\hat{g}_t^i\right] \\
& + m_{3,t}\left[\pi_t^i - \pi_t^* + s_t^i - s_{t-1}^i\right].
\end{aligned}
$$

Note that while $E_t\{\pi_{t+1}^i\}$ is taken as given, it is a given function of today's state variable s_t^i. The first order conditions are given by:

$$
\frac{\partial \mathcal{L}_t}{\partial \pi_t^i} = -\frac{\varepsilon}{\lambda}\pi_t^{i,N} + m_{2,t} + m_{3,t} = 0 \tag{1.E.1}
$$

$$
\frac{\partial \mathcal{L}_t}{\partial \hat{y}_t^i} = -(1+\varphi)\hat{y}_t^{i,N} + \frac{\gamma^N}{1-\gamma^N}(\hat{g}_t^{i,N} - \hat{y}_t^{i,N}) + m_{1,t} - \lambda\left(\frac{1}{1-\gamma^N} + \varphi\right)m_{2,t} = 0 \tag{1.E.2}
$$

$$
\frac{\partial \mathcal{L}_t}{\partial \hat{g}_t^i} = -\frac{\gamma^N}{1-\gamma^N}(\hat{g}_t^{i,N} - \hat{y}_t^{i,N}) - \gamma^N m_{1,t} + \frac{\lambda\gamma^N}{1-\gamma^N}m_{2,t} = 0 \tag{1.E.3}
$$

$$
\frac{\partial \mathcal{L}_t}{\partial s_t^i} = \beta E_t \frac{\partial V(s_t^{i,N}, \pi_{t+1}^{*,N}, \hat{y}_{t+1}^{*,N}, \hat{g}_{t+1}^{*,N})}{\partial s_t^{i,N}} - (1-\gamma^N)m_{1,t} - m_{2,t}\frac{\beta \partial E_t\{\pi_{t+1}^{i,N}\}}{\partial s_t^{i,N}} + m_{3,t} = 0. \tag{1.E.4}
$$

Solving first order conditions (1.E.2) and (1.E.3) for $m_{2,t}$ yields

$$
m_{2,t} = -\frac{1}{\lambda}\hat{y}_t^{i,N} - \frac{1}{\lambda\varphi}\hat{g}_t^{i,N}.
$$

This implies for $m_{3,t}$ by (1.E.1)

$$
m_{3t} = \frac{\varepsilon}{\lambda}\pi_t^{i,N} + \frac{1}{\lambda}\hat{y}_t^{i,N} + \frac{1}{\lambda\varphi}\hat{g}_t^{i,N}.
$$

First order condition (1.E.3) requires for $m_{1,t}$ that

$$m_{1,t} = -\frac{1}{1-\gamma^N}(\hat{g}_t^{i,N} - \hat{y}_t^{i,N}) - \frac{1}{1-\gamma^N}\hat{y}_t^{i,N} - \frac{1}{1-\gamma^N}\frac{1}{\varphi}\hat{g}_t^{i,N}.$$

Finally, substituting for the multipliers in (1.E.4) gives

$$-\lambda\beta E_t \frac{\partial V(s_t^{i,N}, \pi_{t+1}^{*,N}, \hat{y}_{t+1}^{*,N}, \hat{g}_{t+1}^{*,N})}{\partial s_t^{i,N}} + \beta\left(\hat{y}_t^{i,N} + \frac{1}{\varphi}\hat{g}_t^{i,N}\right)\frac{\partial E_t\{\pi_{t+1}^{i,N}\}}{\partial s_t^{i,N}} + \pi_t^{i,N} + \frac{1}{\varepsilon}\hat{y}_t^{i,N} = -\psi_g^N \hat{g}_t^{i,N},$$

with $\psi_g^N \equiv \frac{1}{\varepsilon\varphi}(\lambda\varphi + (1+\lambda))$. Taking the limit of $\beta \to 0$ the above condition becomes

$$\pi_t^{i,N} + \frac{1}{\varepsilon}\hat{y}_t^{i,N} = -\psi_g^N \hat{g}_t^{i,N}. \tag{1.E.5}$$

1.F Proof of proposition 2

In proposition 2 we state the solution for optimal government consumption at the ELB in the absence of coordination when $\beta \to 0$ and compare it to the case of coordination. Under the ELB scenario under consideration (see Section 1.2.2), we get a two-state solution in a symmetric Nash equilibrium. Optimal policy at the ELB is determined by equations (1.14) and (1.16) with $\beta \to 0$ and (1.E.5). In a symmetric equilibrium the resulting system of equations is isomorphic to the one obtained under coordination in Appendix 1.C. However, the definitions of γ^N, κ^N and ψ_g^N differ and now we let $\beta \to 0$. The optimal policy is given by:

$$\hat{g}_L^{*,N} = -\Theta_{\beta\to 0}^N r_L,$$

with

$$\Theta_{\beta\to 0}^N \equiv \frac{(1-\gamma^N)\left(\kappa^N\varepsilon + 1\right)}{\varepsilon\psi_g^N\Gamma_1^N + (1-\mu)\gamma^N\lambda\varphi\varepsilon + \gamma^N\Gamma_2^N},$$

where $\Gamma_1^N \equiv (1-\mu) - (1-\gamma^N)\mu\kappa^N$ and $\Gamma_2^N \equiv (1-\mu) - \mu\lambda$. Since we consider only uniquely determined equilibria (see the ELB scenario in Section 1.2.2), Γ_1^N and Γ_2^N are positive. All other expressions in $\Theta_{\beta\to 0}^N$ are non-negative. Hence, $\Theta_{\beta\to 0}^N > 0$.

Next we show that $\Theta_{\beta\to 0}^N < \Theta_{\beta\to 0}^C$ where Θ^C is as stated in Proposition 1. Since we showed that all terms in $\Theta_{\beta\to 0}^N$ and $\Theta_{\beta\to 0}^C$ are non-negative, we prove the above inequality by showing that the following holds:

$$\frac{\psi_g^C \varepsilon \Gamma_{1,\beta\to 0}^C + (1-\mu)\gamma^C \lambda\varphi\varepsilon + \gamma^C \Gamma_{2,\beta\to 0}^C}{\psi_g^N \varepsilon \Gamma_1^N + (1-\mu)\gamma^N \lambda\varphi\varepsilon + \gamma^N \Gamma_2^N} < 1 < \frac{(1-\gamma^C)\left(\kappa^C\varepsilon + 1\right)}{(1-\gamma^N)\left(\kappa^N\varepsilon + 1\right)}, \tag{1.F.1}$$

where Γ_1^C and Γ_2^C are as stated in Proposition 1.

The left hand side of equation (1.F.1) can be rearranged to:

$$0 < \frac{1}{\varphi}\lambda(1+\varphi)\Gamma_1^N + (\gamma^N - \gamma^C)(1-\mu)(\lambda\varphi\varepsilon + 1),$$

which holds true since $\gamma^N > \gamma^C$ while the remaining terms are positive. Hence, we have established that the left hand side is below unity.

We continue with the right hand side of (1.F.1) which can be rearranged to:

$$(\gamma^N - \gamma^C)(\lambda\varphi\varepsilon + 1) > 0,$$

which holds true since $\gamma^N > \gamma^C$ while the remaining parameters are positive. Hence, we have established that the right hand side is above unity. ■

1.G Numerical solution in the absence of coordination

In order to solve the model numerically we use the algorithm put forward by Soderlind (1999). For this purpose we cast the equilibrium conditions in the following form:

$$E_t \begin{bmatrix} x_{1t+1} \\ x_{2t+1} \end{bmatrix} = A \begin{bmatrix} x_{1t} \\ x_{2t} \end{bmatrix} + Bu_t. \tag{1.G.1}$$

In the expression above x_{1t} are state variables and x_{2t} are control variables. For notational convenience we define $x_t \equiv [x_{1t}, x_{2t}]'$. The policy instrument is denoted by u_t. We rearrange the equilibrium condition at the country level—that is equations (1.7), (1.10) and (1.12) all approximated around the steady state in the absence of coordination—in the following way in order to cast them into (1.G.1):

$$
\begin{aligned}
E_t\{\pi_{t+1}^i\} =& \frac{1}{\beta}\left(1+\kappa^N(1-\gamma^N)\right)\pi_t^i + \frac{\gamma^N}{\beta}\left(\frac{\lambda}{1-\gamma^N}-\kappa^N\right)\hat{g}_t^i - \frac{1}{\beta}\kappa^N\hat{y}_t^* + \frac{1}{\beta}\kappa^N\gamma^N\hat{g}_t^* \\
& - \frac{1}{\beta}\kappa^N(1-\gamma^N)\pi_t^* - \frac{1}{\beta}\kappa^N(1-\gamma^N)s_{t-1}^i, \\
s_t^i =& \pi_t^* - \pi_t^i + s_{t-1}^i.
\end{aligned}
$$

The vectors in (1.G.1) are thus given by

$$
x_{1t} = \begin{bmatrix} s_{t-1}^i \\ \pi_t^* \\ \hat{y}_t^* \\ \hat{g}_t^* \end{bmatrix}, \; x_{2t} = \begin{bmatrix} \pi_t^i \end{bmatrix}, \; u_t = \begin{bmatrix} \hat{g}_t^i \end{bmatrix},
$$

while matrices A and B are given by

$$
A = \begin{bmatrix}
1 & 1 & 0 & 0 & -1 \\
0 & \mu & 0 & 0 & 0 \\
0 & 0 & \mu & 0 & 0 \\
0 & 0 & 0 & \mu & 0 \\
-\frac{1}{\beta}\kappa^N(1-\gamma^N) & -\frac{1}{\beta}\kappa^N(1-\gamma^N) & -\frac{1}{\beta}\kappa^N & \frac{1}{\beta}\kappa^N\gamma^N & \frac{1}{\beta}\left(1+\kappa^N(1-\gamma^N)\right)
\end{bmatrix},
$$

$$
B = \begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \\ \frac{\gamma^N}{\beta}\left(\frac{\lambda}{1-\gamma^N}-\kappa^N\right) \end{bmatrix},
$$

where we use that aggregate variables inherit the Markov structure of the shock in the crisis scenario (see Section 1.2.2).

Under discretion the policy problem is given in general terms by the following expression:

$$x_{1t}'Vx_{1t} + v_t = \min_{u_t} \left[x_t'Qx_t + 2x_t'Uu_t + u_t'Ru_t + \beta E_t\{x_{1t+1}'Vx_{1t+1} + v_{t+1}\} \right] \qquad (1.G.2)$$
$$\text{s.t.} \qquad E_t x_{2t+1} = CE_t x_{1t+1}, \text{ Eq. (1.G.1), and } x_{1t} \text{ given,}$$

where $x_{1t}'Vx_{1t}$ is the value function (quadratic in the state variables), $x_t'Qx_t + 2x_t'Uu_t + u_t'Ru_t$ is the period loss function and $x_{2t} = Cx_{1t}$ is a linear function that maps state variables into control variables.

In order to cast the period loss function into this setup we rearrange it to:

$$\frac{\varepsilon}{\lambda}(\pi_t^i)^2 + \left((1+\varphi) + \frac{\gamma^N}{1-\gamma^N} \right) (\hat{y}_t^i)^2 - 2\frac{\gamma^N}{1-\gamma^N}\hat{y}_t^i\hat{g}_t^i + \frac{\gamma^N}{1-\gamma^N}(\hat{g}_t^i)^2.$$

We further define the auxiliary matrix W by rewriting the above equation as follows:

$$\begin{bmatrix} \pi_t^i & \hat{y}_t^i & \hat{g}_t^i \end{bmatrix} \begin{bmatrix} \frac{\varepsilon}{\lambda} & 0 & 0 \\ 0 & (1+\varphi) + \frac{\gamma^N}{1-\gamma^N} & -\frac{\gamma^N}{1-\gamma^N} \\ 0 & -\frac{\gamma^N}{1-\gamma^N} & \frac{\gamma^N}{1-\gamma^N} \end{bmatrix} \begin{bmatrix} \pi_t^i \\ \hat{y}_t^i \\ \hat{g}_t^i \end{bmatrix} = \begin{bmatrix} \pi_t^i & \hat{y}_t^i & \hat{g}_t^i \end{bmatrix} W \begin{bmatrix} \pi_t^i \\ \hat{y}_t^i \\ \hat{g}_t^i \end{bmatrix}.$$

Further, we use the following version of the market clearing condition at the country level

$$\hat{y}_t^i = \hat{y}_t^* + \gamma^N \hat{g}_t^i - \gamma^N \hat{g}_t^* + (1-\gamma^N)(\pi_t^* - \pi_t^i + s_{t-1}^i)$$

to define the auxiliary matrix K:

$$\begin{bmatrix} \pi_t^i \\ \hat{y}_t^i \\ \hat{g}_t^i \end{bmatrix} = \begin{bmatrix} 0 & 0 & 0 & 0 & 1 & 0 \\ (1-\gamma^N) & (1-\gamma^N) & 1 & -\gamma^N & -(1-\gamma^N) & \gamma^N \\ 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} s_{t-1}^i \\ \pi_t^* \\ \hat{y}_t^* \\ \hat{g}_t^* \\ \pi_t^i \\ \hat{g}_t^i \end{bmatrix} = K \begin{bmatrix} x_t \\ u_t \end{bmatrix}.$$

Then, the period loss function can be cast into (1.G.2) as follows:

$$\begin{bmatrix}\pi_t^i & \hat{y}_t^i & \hat{g}_t^i\end{bmatrix}\begin{bmatrix}\frac{\varepsilon}{\lambda} & 0 & 0\\ 0 & (1+\varphi)+\frac{\gamma^N}{1-\gamma^N} & -\frac{\gamma^N}{1-\gamma^N}\\ 0 & -\frac{\gamma^N}{1-\gamma^N} & \frac{\gamma^N}{1-\gamma^N}\end{bmatrix}\begin{bmatrix}\pi_t^i\\ \hat{y}_t^i\\ \hat{g}_t^i\end{bmatrix} = \begin{bmatrix}x_t' & u_t'\end{bmatrix} K'WK \begin{bmatrix}x_t\\ u_t\end{bmatrix}.$$

By appropriate partitioning of $K'WK$ we obtain the matrices in (1.G.2):

$$\begin{bmatrix}Q & U\\ U' & R\end{bmatrix} = K'WK.$$

The solution to (1.G.2) gives the following policy rule (see Soderlind, 1999):

$$u_t = -Fx_{1t}.$$

Put differently

$$\hat{g}_t^i = -f_1 s_{t-1}^i - f_2\pi_t^* - f_3\hat{y}_t^* - f_4\hat{g}_t^*.$$

In a symmetric equilibrium the terms of trade are zero and the equilibrium is determined at the union level. The equilibrium conditions in the ELB scenario are:

$$\begin{aligned}
\hat{g}_L^{*,N} &= -f_2\pi_L^{*,N} - f_3\hat{y}_L^{*,N} - f_4\hat{g}_L^{*,N}\\
\pi_L^{*,N} &= \frac{1}{1-\beta\mu}\kappa^N(\hat{y}_L^{*,N} - \frac{\bar{\sigma}\gamma^N}{\bar{\sigma}+\varphi}\hat{g}_L^{*,N})\\
\hat{y}_L^{*,N} &= \frac{(1-\gamma^N)(1-\beta\mu)}{(1-\mu)(1-\beta\mu)-(1-\gamma^N)\mu\kappa^N}r_L + \frac{(1-\mu)(1-\beta\mu)\gamma^N - (1-\gamma^N)\mu\kappa^N\frac{\gamma^N\bar{\sigma}}{\bar{\sigma}+\varphi}}{(1-\mu)(1-\beta\mu)-(1-\gamma^N)\mu\kappa^N}\hat{g}_L^{*,N}.
\end{aligned}$$

Given the numerical solution for f_2, f_3 and f_4 we solve the above system for $\hat{g}_L^{*,N}$.

1.H Optimal policy with coordination and commitment

Under commitment we assume that the conduct of monetary policy is described by the following version of the Taylor rule

$$i_t^* = \psi_t\left(r_t + \phi_\pi\pi_t^*\right),$$

where ψ_t is a regime-switching parameter with

$$\psi_t = \begin{cases} 0 & \text{if } r_t < 0, \\ 1 & \text{otherwise.} \end{cases}$$

The Lagrangian is given by

$$\begin{aligned}
\mathcal{L}_t = -\frac{1}{2} E_t \sum_{t=0}^{\infty} \beta^t \Bigg\{ & \left(\frac{\varepsilon}{\lambda} (\pi_t^*)^2 + (1+\varphi)(\hat{y}_t^*)^2 + \frac{\gamma^C}{1-\gamma^C} (\hat{g}_t^* - \hat{y}_t^*)^2 \right) \\
& + \eta_{0,t}^* \left[\beta E_t\{\pi_{t+1}^*\} + \kappa^C \hat{y}_t^* - \frac{\lambda \gamma^C}{1-\gamma^C} \hat{g}_t^* - \pi_t^* \right] \\
& + \eta_{1,t}^* \Bigg[E_t\{\hat{y}_{t+1}^*\} - \hat{y}_t^* - (1-\gamma^C)(\psi_t(r_t + \phi_\pi \pi_t^*) - E_t\{\pi_{t+1}^*\} - r_t) \\
& \qquad - \gamma^C E_t\{\hat{g}_{t+1}^*\} + \gamma^C \hat{g}_t^* \Bigg] \Bigg\}.
\end{aligned}$$

Assuming policymaking under a timeless perspective, the first order conditions read as follows:

$$\frac{\partial \mathcal{L}_t}{\partial \pi_t^*} = -\frac{\varepsilon}{\lambda} \pi_t^{*,C} - \eta_{0,t}^* + \eta_{0,t-1}^* - (1-\gamma^C)\phi_\pi \psi_t \eta_{1,t}^* + \beta^{-1}(1-\gamma^C)\eta_{1,t-1}^* = 0 \tag{1.H.1}$$

$$\frac{\partial \mathcal{L}_t}{\partial \hat{y}_t^*} = -(1+\varphi)\hat{y}_t^{*,C} + \frac{\gamma^C}{1-\gamma^C}(\hat{g}_t^{*,C} - \hat{y}_t^{*,C}) + \kappa^C \eta_{0,t}^* + \beta^{-1}\eta_{1,t-1}^* - \eta_{1,t}^* = 0 \tag{1.H.2}$$

$$\frac{\partial \mathcal{L}_t}{\partial \hat{g}_t^*} = -\frac{\gamma^C}{1-\gamma^C}(\hat{g}_t^{*,C} - \hat{y}_t^{*,C}) - \frac{\lambda\gamma^C}{1-\gamma^C}\eta_{0,t}^* - \beta^{-1}\gamma^C \eta_{1,t-1}^* + \gamma^C \eta_{1,t}^* = 0 \tag{1.H.3}$$

The equilibrium conditions under coordination with commitment are thus given by equations (1.8), (1.11), (1.H.1), (1.H.2) and (1.H.3). We solve for the equilibrium as described in Appendix A in Eggertsson and Woodford (2003). They compute optimal monetary policy in a similar ELB scenario. In order to solve for optimal policy under commitment they assume that there is a certain date S at which the shock to r_t disappears with probability one. Given the solution for this particular case, one can then compute the solution for any period $t < S$ in which the shock disappears unexpectedly. To mimic our ELB scenario in Section 1.2.2, we assume that period S is sufficiently far in the future. Furthermore, compared to their analysis, the Taylor rule in our setup implies that there is no state in which the ELB is still binding even though the shock disappeared.

1.I Consumption equivalent

We compute the consumption equivalent ζ as the compensation in terms of consumption which is required to make the household indifferent between the optimal stimulus under coordination and the stimulus which is optimal in the absence of coordination. Importantly, we compare the fiscal stimuli with and without coordination around the same steady state—namely the one with coordination. We use superscript $\tilde{N}$ to index the equilibrium outcome which obtains if the percentage increase of government spending at the efficient steady state is equal to what is optimal in the absence of coordination. We use ζ to denote the percentage increase of consumption that makes the household in this equilibrium as well off as in the equilibrium with the stimulus under coordination (as given in proposition 1). Formally, the consumption equivalent is defined by

$$U_L^C = U_L^{\tilde{N}}(\zeta), \tag{1.I.1}$$

where

$$\begin{aligned} U_L^C &= (1-\chi)\log C_L^C + \chi \log G_L^C - \frac{\left(N_L^C\right)^{1+\varphi}}{1+\varphi}, \\ U_L^{\tilde{N}}(\zeta) &\equiv (1-\chi)\log(C_L^{\tilde{N}}(1+\zeta)) + \chi \log G_L^{\tilde{N}} - \frac{\left(N_L^{\tilde{N}}\right)^{1+\varphi}}{1+\varphi} \\ &= U_L^{\tilde{N}}(0) + (1-\chi)\log(1+\zeta). \end{aligned}$$

Here, we assume that the household receives the compensation only for as long as the ELB binds. Inserting in (1.I.1) and rearranging yields:

$$(1-\chi)\log(1+\zeta) = U_L^C - U_L^{\tilde{N}}(0).$$

We use the second order approximation to period utility provided by Galí and Monacelli (2008) to approximate U_L^C and $U_L^{\tilde{N}}(0)$ around the steady state under coordination. After rearranging we get:

$$\zeta \approx \exp\Bigg\{ -\frac{1}{2}\frac{1}{1-\chi}\left(\frac{\varepsilon}{\lambda}(\pi_L^{*,C})^2 + (1+\varphi)(\hat{y}_L^{*,C})^2 + \frac{\gamma^C}{1-\gamma^C}(\hat{g}_L^{*,C} - \hat{y}_L^{*,C})^2\right) + \frac{1}{2}\frac{1}{1-\chi}\left(\frac{\varepsilon}{\lambda}(\pi_L^{*,\tilde{N}})^2 + (1+\varphi)(\hat{y}_L^{*,\tilde{N}})^2 + \frac{\gamma^C}{1-\gamma^C}(\hat{g}_L^{*,\tilde{N}} - \hat{y}_L^{*,\tilde{N}})^2\right)\Bigg\} - 1. \tag{1.I.2}$$

Chapter 2

Exchange Rate Undershooting: Evidence and Theory

Joint with Gernot Müller and Martin Wolf

2.1 Introduction

Some 40 years ago, Dornbusch (1976) put forward a seminal account of how exchange rates adjust to monetary policy shocks. It goes as follows. In the long run the exchange rate appreciates in response to a contractionary monetary policy shock. This ensures that purchasing power parity will be restored since the shock induces a permanent decline of the price level. In the short run, as domestic interest rates exceed foreign rates, market participants expect the exchange rate to depreciate. This ensures that uncovered interest parity will be satisfied. How can expectations of an appreciation in the long run be consistent with expectations of a depreciation in the short-run? The exchange rate has to overshoot its new long run level on impact.

Expectations take center stage in Dornbusch's account and, importantly, adjust instantaneously to the shock. Yet by now there is pervasive evidence that expectations adjust only sluggishly to new information (Coibion and Gorodnichenko, 2012, 2015). More specifically, recent work by Nakamura and Steinsson (2018a) illustrates the importance of information frictions for the transmission of monetary policy.[1] Since market participants have incomplete information about the state of the economy, monetary policy innovations carry potentially new information about the natural rate of interest.

This matters for exchange rate dynamics because natural rate shocks that signal rising potential output will generally depreciate the exchange rate. Due to information frictions

[1]Romer and Romer (2000) established the notion that there are informational asymmetries between the Federal Reserve and the private forecasters. They also show that private forecasters adjust forecasts in response to monetary policy actions. Melosi (2017) estimates a model with dispersed information. In his model monetary policy shocks have a "signaling effect" to the extent that heterogenous price setters seek to learn about the state of the economy.

market participants only learn over time the true nature of the shock. As a result, the initial response of the nominal exchange rate to a genuine monetary policy shock may therefore be muted. The exchange rate, in other words, undershoots rather than overshoots its new long-run value in the presence of information frictions. In the first part of this paper we provide evidence that the exchange rate indeed undershoots in response to monetary policy shocks. In the second part of the paper we put forward an open economy model with information frictions. It is able to account for the evidence, both qualitatively and quantitatively.

In the first part of the paper, we rely on local projections to estimate the response of the exchange rate and of other variables of interest to monetary policy shocks (Jordà, 2005). For this purpose we use the series of monetary policy shocks identified by Romer and Romer (2004) and updated by Coibion et al. (2017). Importantly, this series represents genuine monetary policy shocks, that is, innovations to the federal funds rate which are purged of not only of endogenous, but also of anticipatory components. In addition, local projections offer a convenient way to capture the permanent effect of monetary policy shocks on the exchange rate.

The sample for our baseline specification runs from the post-Bretton Woods period to 2008 and we obtain estimates based, in turn, on quarterly and monthly data. Our main variable of interest is the effective exchange rate of the US dollar, but we also estimate the effect of US monetary policy shocks on US output and prices. The key result of our empirical analysis is that the exchange rate undershoots its new long-run level. In response to an exogenous interest rate increase by 100 basis points, the exchange rate appreciates by about 1 percent on impact. Over time there is further appreciation with a maximum effect of about 5 percent after about 2-3 years.

We seek to rationalize the evidence in the second part of the paper as we develop a New Keynesian model in which agents form expectations rationally, but subject to an information friction. In particular, whereas the central bank has full knowledge of the natural rate, private agents observe the natural rate imperfectly, as in Nakamura and Steinsson (2018a). Market participants therefore face an inference problem whenever they observe a surprise increase of the interest rate. It may reflect a genuine monetary policy shock, a policy response to a natural rate shock, or, as in Nakamura and Steinsson's formulation, a mixture of both. As a consequence, interest rate surprises induce an "information effect" that generates an adjustment that differs from what obtains in full-information models.

Our model differs from the model of Nakamura and Steinsson (2018a) in two respects.

First, we consider a small open economy rather than a closed economy model. Second, our model features both a genuine monetary policy shock and a natural rate shock. As a result, we have to make the inference problem of market participants explicit. At the same time, this allows us to study the effect of monetary policy shocks and natural rate shocks in isolation—both under perfect and imperfect information. Under perfect information, a monetary policy shock generates nominal appreciation and exchange rate overshooting, just like in Dornbusch (1976). Following a natural rate shock, on the other hand, the exchange rate *depreciates* and more so in the long run than in the short run because the natural rate shock signals a rising level of potential output.

The key finding of our model analysis is that a monetary policy shock may induce exchange rate undershooting in the presence of information frictions. Intuitively, market participants account for the possibility that the policy rate increase is due to an increase of the natural rate, in which case the exchange rate would depreciate in the long run. As a forward looking variable, the immediate response of the exchange rate reflects this possibility as well as the possibility that the increase of the policy rate represents a genuine shock. Its response is therefore dampened in the short run, both relative to a full-information scenario and relative to the long-run response. It is only over time that agents learn the true nature of the shock and the exchange rate continues to appreciate.

We also perform a quantitative analysis and estimate key parameters of the model by matching the empirical impulse response functions for the federal funds rate, output, the price level and the nominal exchange rate. Specifically, we employ an indirect inference approach using the estimated impulse response functions as an "identified moment" that conveys actual information about structural features of the economy (Nakamura and Steinsson, 2018b). We use this procedure for it is quite robust to misspecification in both the empirical model that is used to generate the identified moment (Gourieroux et al., 1993; Smith, 2008), as well as in the theoretical model that is used to explain the identified moment (Nakamura and Steinsson, 2018b). As explained in Smith (2008), under some conditions, the estimated parameters are the same as those obtained by maximum likelihood.

We find that the estimated model predicts a path for the exchange rate in response to a monetary policy shock which aligns closely with the empirical impulse response function for this variable. The estimated parameters also appear plausible. In particular, they imply that market participants attribute about 2/3 of the innovations of the policy rate to natural rate disturbances, and only 1/3 to genuine monetary policy shocks. This value is almost identical to the estimate reported by Nakamura and Steinsson (2018a) on the basis of an altogether different approach. Also, the extent of information friction implied by our estimates squares well with the results of Coibion and Gorodnichenko (2012).

In an influential study, Eichenbaum and Evans (1995) estimated the effects of monetary policy shocks within a vector autoregression (VAR) model and documented a pattern that has become known as "delayed overshooting": in response to a contractionary monetary shock the exchange rate appreciates on impact and depreciates thereafter. However, the depreciation sets in only after a delay of more than two years. During the interim period, the exchange rate continues to appreciate. This pattern has been found to be robust across a number of alternative specifications and identification schemes. It has also been documented for the real exchange rate (see, for instance, Scholl and Uhlig, 2008; Steinsson, 2008; Bouakez and Normandin, 2010; Bruno and Shin, 2015).[2] In a recent contribution, Kim et al. (2017), have performed a sub-sample analysis and found that delayed overshooting obtains only during the 1980s.

We also find for some specifications of our empirical model a reversal of exchange rate dynamics after the maximum effect—in line with delayed overshooting. However, regarding this reversal, we find a) that it occurs very late (after about 4 years) and b) that it is rather modest. Also, in a recent paper, Schmitt-Grohé and Uribe (2018) distinguish between temporary and permanent shocks as they identify the effect of monetary policy on the exchange rate in an estimated state-space model. They find no overshooting for either temporary nor permanent shocks—neither immediate, nor delayed. Against this background, we consider "undershooting" rather than "delayed overshooting" the core issue on which we focus in our analysis.

Still, we acknowledge related work that has rationalized delayed overshooting on theoretical grounds. In Bacchetta and van Wincoop (2010) delayed overshooting arises because of infrequent portfolio adjustments. Kim et al. (2017) argue that delayed overshooting arises in the 1980s because the Volker disinflation was lacking credibility. This is consistent with the model put forward by Gourinchas and Tornell (2004) in which investors systematically underestimate the persistence of interest rate innovations. Importantly, in our model we rely on a different kind of information friction. We assume that private agents cannot distinguish monetary shocks from natural rate surprises. And while we show that our estimated model accounts for undershooting, we stress that it is also able to generate either

[2] However, the "delayed overshooting puzzle" has not gone unchallenged and some studies have indeed reported overshooting (Kim and Roubini, 2000; Faust and Rogers, 2003; Bjornland, 2009; Forni and Gambetti, 2010).

immediate overshooting or delayed overshooting for alternative parameterizations.

The remainder of the paper is organized as follows. In the next section we describe our empirical analysis and present results. Section 2.3 outlines the small open economy model with information rigidities. It also explains our estimation procedure and discusses results. Section 2.4 inspects the mechanism, notably by contrasting the transmission mechanism in the model to the one in which market participants have full information. In Section 2.5 we consider additional evidence which has not been used in the estimation to assess the external validity of our results. A final section concludes.

2.2 Evidence

In this section we provide evidence on how the nominal exchange rate adjusts to monetary policy shocks. We focus on the US and estimate the response of the effective exchange rate of the US dollar to a shock in the federal funds rate. We use series of monetary policy shocks provided by Coibion et al. (2017) who, in turn, extend and update the original shock series identified by Romer and Romer (2004). We use the narrow nominal effective exchange rate index compiled the Bank for International Settlement (BIS) which is a trade-weighted index of bilateral exchange rates of the US to 14 economies and the euro area.

The approach to identify the shocks is detailed in Romer and Romer (2004). Here we summarize the main idea. In a first step, Romer and Romer construct series for the Fed's intended federal funds rate before FOMC meetings on the basis of narrative sources. In a second step Romer and Romer purge the changes of the intended federal funds rate at FOMC meetings from changes that are due to information the Fed has about the future economic development. For this purpose they regress the change of the intended federal funds rate on the Fed's Greenbook forecasts for inflation, real output growth, and the unemployment rate. Since there is no evidence that monetary policymakers use a substantial amount of information on economic activity in addition to the Greenbook forecasts as they set the rate, the regression residuals may be interpreted as genuine monetary policy shocks.

We use local projections to directly estimate the impulse responses to monetary policy shocks (Jordà, 2005). For our baseline specification, we estimate the responses of the nominal exchange rate and of key macroeconomic variables using quarterly data for the period 1976Q1 until 2007Q3, that is, our sample starts after the Bretton Woods System had been completely abandoned (see also Kim et al., 2017). Our sample stops before the financial crisis. In our robustness analysis we consider alternative sample periods as well.

Our empirical specification builds on Coibion et al. (2017). Formally, letting e_t^{RR} denote a US monetary policy shock in period t and x_t the realization of a variable of interest, we estimate the following model:

$$x_{t+h} - x_{t-1} = c^{(h)} + \sum_{j=1}^{J} \alpha_j^{(h)} (x_{t-j} - x_{t-j-1}) + \sum_{k=0}^{K-1} \beta_k^{(h)} e_{t-k}^{RR} + \varepsilon_{t+h}. \tag{2.1}$$

In this specification, we estimate the effect on the variable of interest at horizon h relative to to the pre-shock level (Stock and Watson, 2018). In this way, we account for permanent effects of monetary policy shocks on the variables of interest. Our specification includes J lags of the dependent variable and K lags of the shock. $c^{(h)}$ is a constant for horizon h and ε_{t+h} is an error term with mean zero and strictly positive variance. We compute heteroscedasticity and autocorrelation consistent standard-errors as in Newey and West (1987).

We estimate the empirical model above for the log of US real GDP, the log of the CPI, as well as for the log of the nominal effective exchange rate of the US dollar. We also estimate the response of the federal funds rate to the shock in levels.[3] We follow Coibion et al. (2017) and restrict the contemporaneous effect of monetary policy shocks on GDP and the CPI to be equal to zero.[4] For all regressions we consider one year of lags of the shock and the endogenous variable, that is, we set $J = 4$ and $K = 4$. Our results are robust across alternative specifications, as we show below.

Figure 2.1 shows the impulse responses to a monetary policy shock which is normalized so that the federal funds rate increases by 100 basis points initially. Here the solid lines represent the point estimate, while shaded areas indicate 68 percent and 90 percent confidence bands. The horizontal axis measures time in quarters. The vertical axis measures the impulse response in percentage points (fed funds rate) or the deviation from the pre-shock level in percent (the other variables). The upper-left panel shows that the federal funds rate rises persistently for about 1.5 years. Afterwards it gradually converges back to zero. The upper right panel shows the response of output which displays a distinct hump-shaped pattern, familiar from earlier work on the monetary transmission mechanism. We observe a maximum effect after about one year, when output has declined by approximately 1 percent. The effect on output ceases to be significant after 2-3 years. The lower-left panel shows

[3] In this case we exclude the pre-shock level and do not consider differenced lags. All series except for the exchange rate (BIS) are obtained from the St. Louis Fed (FRED).

[4] In this case, the second sum in equation (2.1) runs from $k = 1$ to K.

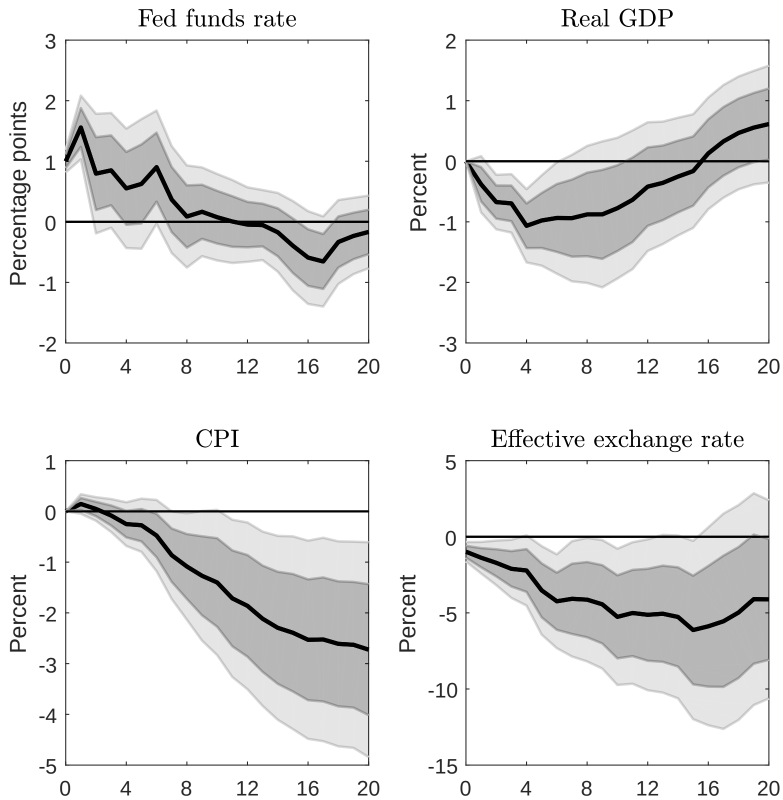

Figure 2.1: Impulse responses to a monetary policy shock. Solid line corresponds to point estimate. Shaded areas correspond to 68% and 90% confidence intervals, respectively. Time (horizontal axis) is in quarters. Vertical axis measures deviation from pre-shock level in percent, except for the federal funds rate (percentage points, annualized). Sample: 1976Q1 to 2007Q3.

the response of the price level. Initially, prices adjust sluggishly. We observe a significant decline of prices only after about 1.5 to 2 years, again a familiar finding of earlier studies. However, the price level continues to decline markedly afterwards. Five years after the shock it is reduced by some 3 percent.

Finally, we turn to response of the nominal exchange rate, shown in the lower-right panel. Here, the exchange rate measures the price of foreign currency in terms of US dollars. Hence, a decline of the exchange rate represents an appreciation of the dollar. We observe a significant impact response. The dollar appreciates immediately by approximately 1 percent in response to the shock. The appreciation, however, continues over time. Only after about two years does exchange rate settle on a new level. At this point is has gained some 5

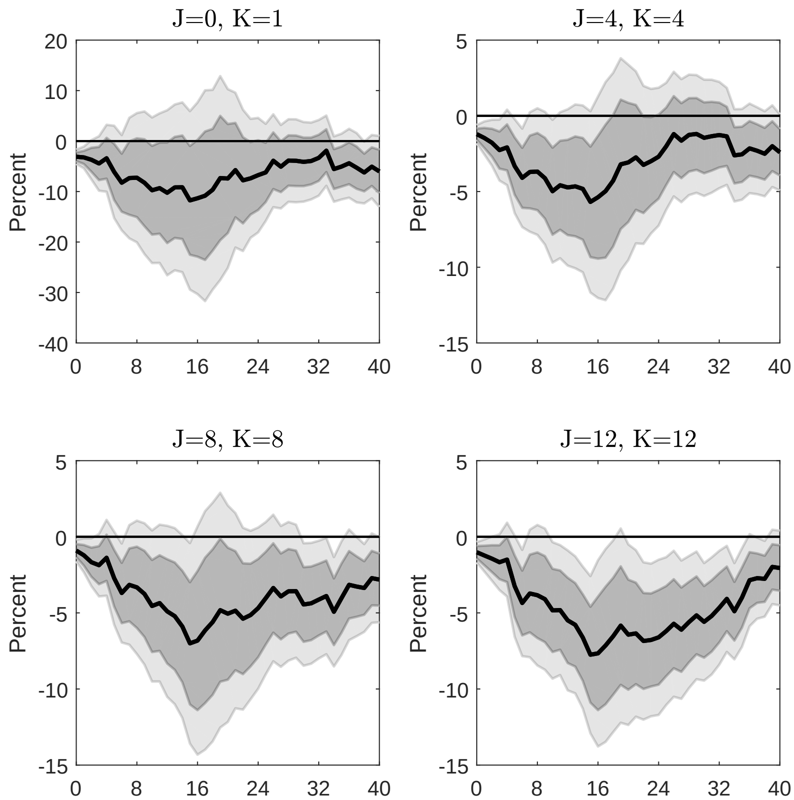

Figure 2.2: Impulse response of the nominal exchange rate for alternative model specifications. J and K refer to number of lags of the dependent variable and the shock included in local projection (2.1), respectively. Solid line corresponds to point estimate. Shaded areas correspond to 68% and 90% confidence intervals, respectively. Time (horizontal axis) is in quarters. Vertical axis measures deviation from pre-shock level in percent (annualized). Sample: 1976Q1 to 2007Q3.

percent in value. Relative to this long-run effect, the impact response is muted. We find, in other words, that the exchange rate undershoots its new long-run level on impact.

In Figure 2.2, we further explore the long-run response of the nominal exchange rate. For this purpose we estimate impulse responses for a horizon of 10 years ($h = 40$). In addition, we consider different specifications for the number of lags of the shock and the nominal exchange rate in the empirical model (2.1). In the upper-left panel we only include the contemporaneous shock. The upper-right panel is our baseline scenario but now estimated for an horizon for up to 10 years. The lower panels increase the number of lags of both the monetary policy shock and the nominal exchange rate to 8 and 12, respectively.

Across specifications the exchange rate dynamics display some notable patterns. First, in response to a monetary contraction the dollar appreciates permanently. Second, the initial response is weaker than the long-run response. In this sense, there is always undershooting. Third, there is some reversal of the maximum appreciation down the road. In principle, this finding is consistent with "delayed overshooting", the focus of much of the earlier literature following Eichenbaum and Evans (1995). However, in what follows we focus on the fact that the exchange rate undershoots initially because the reversal of the exchange rate occurs very late and is fairly modest. Also, at no point is this reversal complete.[5]

Next, we estimate model (2.1) on monthly data. In line with the baseline specification, we now include 12 lags of the shock and the dependent variable ($J = 12$ and $K = 12$). In this context we also estimate the response of the narrow real effective exchange rate (also taken from the BIS). Further, we consider alternative horizons in our estimation. Figure 2.3 shows the results. The horizontal axis now measures time in months. The upper panels show the impulse response of the nominal and the real exchange rate for the short run, that is, during the first year after the shock ($h = 12$). Both, the nominal and real exchange rate appreciate significantly at the 90% level on impact by about 0.5 percent and continue to appreciate during the entire year. The value of the dollar increase by some 4 percent, both nominally and in real terms at the end of the first year.

The lower panels of Figure 2.3 show the response of the nominal and the real exchange rate for a horizon of 5 years. The nominal exchange rate continues to appreciate with a maximum effect of about 8 percent after some 4 years. Hence, we find that undershooting is a robust feature of the data. It obtains for monthly data as well. The impulse response of the real exchange rate resembles the one of the nominal exchange rate, a familiar finding from earlier VAR studies (for instance, Eichenbaum and Evans, 1995).

Finally, we consider alternative sample periods. The Romer-Romer shock series compiled by Coibion et al. (2017) ends in 2008. For the more recent period we rely on monetary policy shocks as identified by Jarociński and Karadi (2018). Jarociński and Karadi (2018) rely on high-frequency data around policy announcement to measure interest rate surprises. A key contribution of their paper is to disentangle genuine monetary policy shocks and complementary central bank information that may also raise interest rates. In order to do

[5] Note that model (2.1) accounts for permanent effects of monetary policy shocks on the exchange rate. Earlier VAR based studies instead include the exchange rate in levels. Yet conventional VAR models deliver inconsistent estimates of impulse response functions at longer horizons in the presence of unit roots (Phillips, 1998).

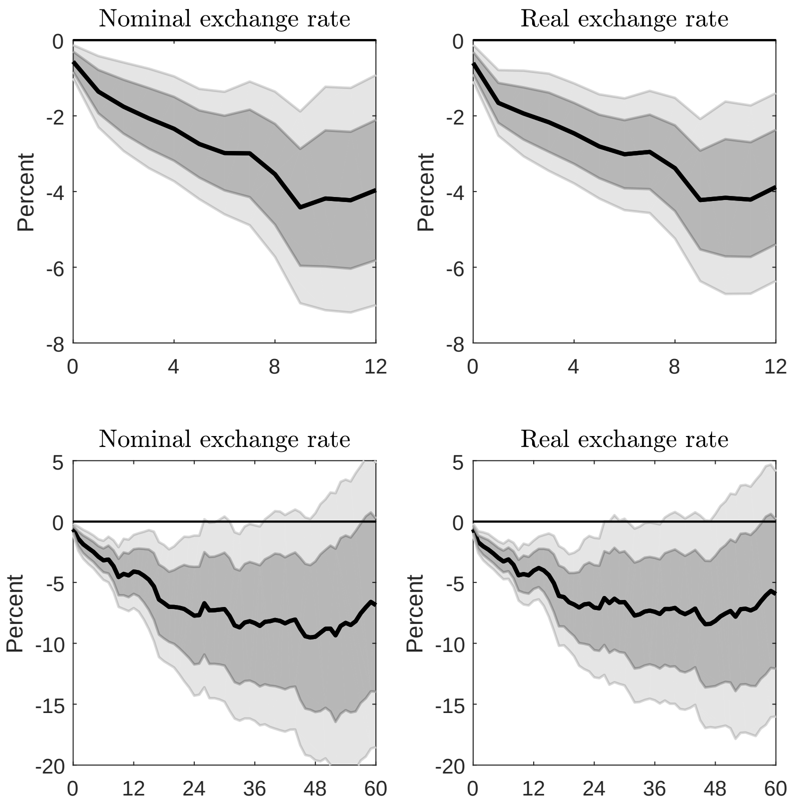

Figure 2.3: Impulse response of the exchange rate to a monetary policy shock: nominal (left) and real (right). Time (horizontal axis) is measured in months. Solid line corresponds to point estimate. Shaded areas correspond to 68% and 90% confidence intervals, respectively. Time (horizontal axis) is in months. Vertical axis measures deviation from pre-shock level in percent annualized). Sample: 1976M1 to 2007M7.

so they restrict the sign of the response of stock prices. These shocks are available for the period between February 1990 and December 2016.

In Figure 2.4 we vary the sample periods on which we estimate our empirical model (2.1). The upper panels first show the results for different samples using the Romer-Romer shocks. In the upper-left panel we consider a longer sample. It starts in 1973 when the Bretton-Woods system broke down and runs until the end of 2008. The results do not change much for this specification relative to the baseline. In the upper-right panel we consider only the post-Volcker period and exclude the financial crisis (1988-01 to 2007-07). For this short sample, we find undershooting as well (although the exchange rate depreciates on impact). This is in contrast to Kim et al. (2017), who find overshooting for this period.

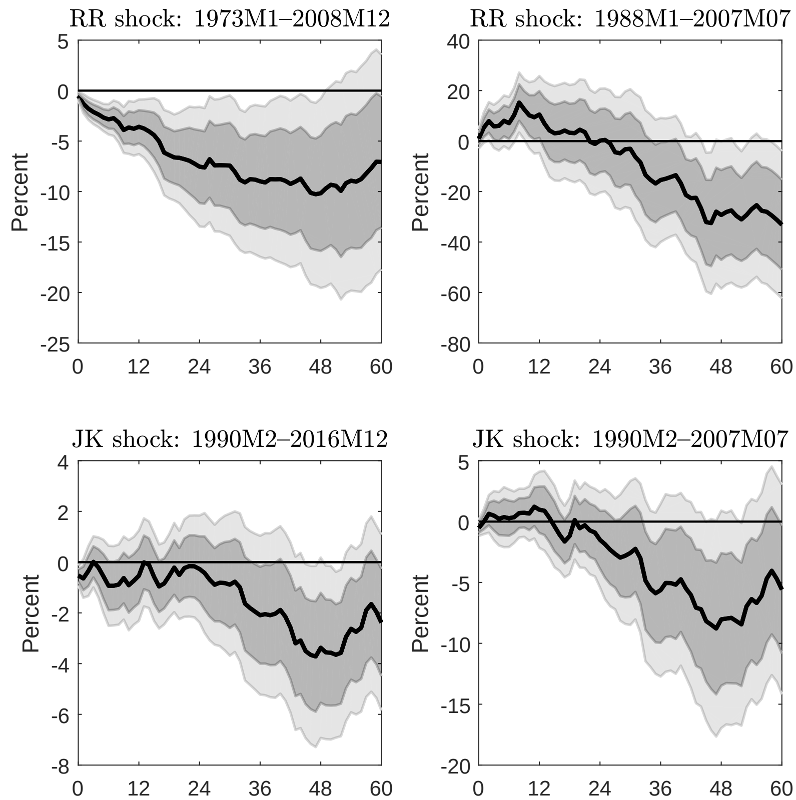

Figure 2.4: Impulse response of the nominal exchange rate to two kinds of monetary policy shocks: Upper panels use Romer, Romer (RR) shocks. Lower panels use Jarociński, Karadi (JK) shocks. Shaded areas correspond to 68% and 90% confidence intervals, respectively. Time (horizontal axis) is in months.

In the lower panels of Figure 2.4, we show results for the Jarociński-Karadi shock series. The lower-left panel shows the impulse response of the nominal exchange rate for the whole sample.[6] The lower-right panel shows results for a shorter sample. It, too, captures the post-Volcker period and runs up until the beginning of the financial crisis. Overall, we find that the pattern of the exchange rate response to the US monetary policy shocks identified by Jarociński and Karadi (2018) is fairly similar to that obtained for the Romer-Romer shocks, even as the sample period differs.

[6]We normalize the impulse response such that the impact equals the one for the Romer-Romer shocks.

2.3 Theory

In this section we put forward our model, which builds on and extends the model by Nakamura and Steinsson (2018a). We estimate the model on the basis of an indirect inference approach. Finally, we use the estimated model to quantify the extent of information frictions that is required to account for the evidence established in Section 2.2.

2.3.1 Model

We consider a New Keynesian small open economy model à la Galí and Monacelli (2005a). The distinct feature of our model are information frictions. The private sector observes only a noisy signal of potential output and therefore of the natural rate of interest. The central bank, in contrast, observes the natural rate perfectly. Market participants understand the central bank's reaction function. As a result, the central bank conveys new information about the natural rate to the private sector whenever it adjusts the policy rate. However, agents update their beliefs about the natural rate only imperfectly, as adjustments in the policy rate may also represent monetary policy shocks.

We connect with two recent advances in the literature on expectations formation and its links with monetary policy. First, we build a model with noisy information. Coibion and Gorodnichenko (2012, 2015) show that information frictions, or more specifically models with noisy information, account well for key features of the data on expectations formation. Second, we build on Nakamura and Steinsson (2018a) who show that a monetary surprise reflects not necessarily a monetary policy shock, but also carries information about the natural rate. We extend the analysis of Nakamura and Steinsson (2018a) in that we i) consider an open economy model, ii) model the process underlying movements in the natural rate explicitly and iii) make the inference problem of agents explicit—that is, we study the case of full rational expectations.[7]

The environment underlying our model is standard except for the information friction. The domestic country is small such that domestic developments have no bearing on the rest of the world. In the domestic economy, monopolistically competitive firms produce a variety of goods which are consumed domestically as well as exported. The law of one

[7] There is an important distinction between expectations not being rational, and expectations being rational but based on incomplete information. Our model belongs to the latter category. The former category includes models where agents form expectations by learning using subjective beliefs (e.g., Adam et al., 2012), or models where agents use adaptive expectations, among many others.

price holds at the level of varieties. Prices are set in the currency of the producer and adjusted infrequently due to a Calvo constraint. Goods markets are imperfectly integrated as domestically produced goods account for a non-zero fraction of the final consumption good. Put differently, the share of domestic goods in home consumption is disproportional to the size of the domestic economy. The real exchange rate may deviate from purchasing power parity as a result. International financial markets are complete so that there is perfect consumption risk sharing between the rest of the world and the domestic country.

In the appendix, we specify the problems of households and firms in detail. In what follows we provide, in turn, a compact exposition of the approximate equilibrium conditions and an explanation of how expectations are formed in the presence of noisy information.

2.3.1.1 Approximate equilibrium conditions

We approximate dynamics in the neighborhood of the steady state. The structural parameters and initial conditions in the domestic economy are the same as in the rest of the world. The steady state is therefore symmetric. There is no inflation in steady state and international relative prices are unity. In what follows, we express all variables in logs. Foreign variables are denoted with a star. They are time-invariant because they are not affected by developments that occur in the (small) domestic economy.

Inflation dynamics are determined by the New Keynesian Phillips curve:

$$\pi_t = \beta \tilde{E}_t \pi_{t+1} + \kappa(y_t - y_t^n) + \eta_t, \tag{2.1}$$

where π_t is inflation of domestically produced goods, y_t is output, y_t^n is potential output, and $\eta_t \sim \mathcal{N}(0, \sigma_\eta^2)$ is an exogenous disturbance. In turn, $0 < \beta < 1$ is the time-discount factor and $\kappa > 0$ captures the extent of nominal rigidities.

The fact that expectations are not based on the full but rather on an incomplete information set is indicated by a tilde above the expectations operator, $\tilde{E}_t$. The way in which $\tilde{E}_t$ is formed and its properties are detailed below.

We assume that potential output follows a first-order autoregressive process in *first differences*

$$\Delta y_t^n = \rho_y \Delta y_{t-1}^n + \varepsilon_t^y, \quad \varepsilon_t^y \sim \mathcal{N}(0, \sigma_y^2),$$

where $0 \leq \rho_y < 1$, such that a positive disturbance $\varepsilon_t^y > 0$ sets in motion a gradual increase of y_t^n to a permanently higher level. This will capture that an increase of the natural rate signals to private agents a growing economy, in line with the evidence put forward by Nakamura and Steinsson (2018a).

A second equilibrium condition links output and the real exchange rate

$$\theta y_t = e_t + p^* - p_t, \tag{2.2}$$

where θ^{-1} is the intertemporal elasticity of substitution. Here, e_t denotes the nominal exchange rate, the price of foreign currency expressed in terms of domestic currency, p_t is the price index of domestically produced goods (such that $\pi_t = p_t - p_{t-1}$) and p^* is the foreign price level. The composite term $e_t + p^* - p_t$ determines the country's terms of trade, which move proportionately with the real exchange rate in our model. An increase in e_t therefore indicates a nominal depreciation of the domestic currency, whereas an increase in $e_t + p^* - p_t$ indicates a *real* depreciation. In order to obtain equation (2.2) we combine market clearing for domestically produced goods with the risk-sharing condition implied by complete international financial markets (Backus and Smith, 1993). Equation (2.2) shows that following a real depreciation the demand for domestically produced goods increases.

The nominal exchange rate, in turn, is determined via the uncovered interest rate parity condition

$$\tilde{E}_t \Delta e_{t+1} = i_t - i^*. \tag{2.3}$$

Here, i_t is the domestic short-term nominal interest rate, and i^* is the foreign rate. According to this condition, the exchange rate is expected to depreciate whenever domestic interest rates exceed foreign rates.

Finally, the model is closed by specifying monetary policy. We posit the following Taylor-type rule

$$i_t = r_t^n + \phi \pi_t + u_t, \tag{2.4}$$

where $\phi > 1$, in line with the Taylor principle, and where u_t is a monetary policy shock, for which we assume

$$u_t = \rho_u u_{t-1} + \varepsilon_t^u, \quad \varepsilon_t^u \sim \mathcal{N}(0, \sigma_u^2).$$

The natural rate, in turn, is defined as the real rate that would prevail absent price and information rigidities. In our model this implies

$$r_t^n \equiv (r_t - E_t\pi_{t+1})|_{\kappa=\infty} = \bar{r} + \theta E_t \Delta y_{t+1}^n = \bar{r} + \theta\rho_y \Delta y_t^n, \tag{2.5}$$

where E_t denotes expectations under full information.[8] Notice that when potential output y_t^n rises, this temporarily *raises* the natural rate, because it foreshadows a growth path along which potential output approaches a permanently higher level.

2.3.1.2 Information processing

We assume that households and firms observe the variables i_t, y_t, e_t and p_t (and therefore π_t and Δe_t) perfectly, whereas the remaining variables u_t, η_t and y_t^n (and therefore r_t^n, by equation (2.5)) are unobserved. However, private agents learn about the unobserved variables as they obtain signals regarding the state of the economy.

More precisely, private agents receive two signals about the unobserved variables. First, because they can observe π_t and y_t perfectly, they can infer the sum $-\kappa y_t^n + \eta_t$ from the Phillips curve (2.1). However, they cannot distinguish the sum's individual components. The disturbance η_t therefore has the natural interpretation of representing noise in the observation of potential output y_t^n; or it may equally be interpreted as any shift in the Phillips curve that is unrelated with changes in potential output, such as cost push shocks, or short-term financing frictions that may affect the domestic firms.

The second signal comes from the Taylor rule (2.4). Because private agents observe i_t and π_t, they can infer the sum $r_t^n + u_t$. Again, they cannot distinguish the sum's individual components. In other words, they cannot tell whether a monetary surprise represents a monetary policy shock u_t or a change in the natural rate r_t^n, in line with the arguments of Nakamura and Steinsson (2018a). In this context we highlight once more that the natural rate r_t^n is tied to the growth rate of potential output according to (2.5). The latter is equally unobserved.

The monetary policy shock u_t thus plays a dual role in our analysis. On the one hand, it is our main object of interest because we study the response of the exchange rate following

[8] To obtain this equation, we set $\kappa = \infty$ in the Phillips curve (2.1) which yields $y_t = y_t^n$. Second, we combine the equation for the real exchange rate (2.2) and the UIP condition (2.3), and replace $y_t = y_t^n$. Finally, we set $i^* = \bar{r} = -\log(\beta) > 0$ because the foreign nominal rate i^* is in its steady state with zero trend inflation. More details on the linearized model can be found in the appendix.

a shock to u_t. However, shocks to u_t are also key for information frictions to impact macroeconomic dynamics. This is because monetary policy shocks provide a *second* source of noise in the observation of the natural rate for private agents. Indeed, in the absence of monetary policy shocks, while private agents could not infer the natural rate from the Phillips curve (2.1), they could do so from the Taylor rule (2.4), and the model would reduce to one of full information.

The key difference between private agents and the central bank in our model is that the central bank can observe the natural rate perfectly. This can be seen by recognizing that the central bank sets its policy rate with reference to the natural rate, in the interest rate feedback rule (2.4). It is not implausible to assume that the central bank has better information about the natural rate than the private sector for two reasons. First, as argued by Nakamura and Steinsson (2018a), it is *optimal* for the central bank in this class of models to set the policy rate with reference to the natural rate. Second, the central bank employs a "legion of PhD economists" who carry out its task to track the natural rate.

At this stage we reemphasize the main difference of our model relative to the model of Nakamura and Steinsson (2018a). In their model, a monetary innovation always represents a *composite* disturbance, as it represents a simultaneous tightening of monetary policy and a rise in the natural rate. In contrast, we construct a model of noisy information in order to separate these two kinds of disturbances: in our setup, shocks to u_t or r_t^n represent independent sources of variation.

Given the above considerations it is straightforward to specify how private agents form expectations. From the previous discussion we have seen that private agents receive two signals about the state of potential output y_t^n and the monetary shock u_t: the sums $-\kappa y_t^n + \eta_t$ and $r_t^n + u_t - \bar{r} = \theta \rho_y \Delta y_t^n + u_t$, where $\Delta y_t^n \equiv y_t^n - y_{t-1}^n$. This is a linear system. Thus, under rational expectations private agents form their beliefs using the *linear Kalman filter*—as in the noisy information models described in Coibion and Gorodnichenko (2012).

Formally, we obtain a state-space representation:

$$\begin{pmatrix} y_t^n \\ y_{t-1}^n \\ u_t \end{pmatrix} = F \begin{pmatrix} y_{t-1}^n \\ y_{t-2}^n \\ u_{t-1} \end{pmatrix} + \begin{pmatrix} \varepsilon_t^y \\ 0 \\ \varepsilon_t^u \end{pmatrix} = \begin{pmatrix} 1+\rho_y & -\rho_y & 0 \\ 1 & 0 & 0 \\ 0 & 0 & \rho_u \end{pmatrix} \begin{pmatrix} y_{t-1}^n \\ y_{t-2}^n \\ u_{t-1} \end{pmatrix} + \begin{pmatrix} \varepsilon_t^y \\ 0 \\ \varepsilon_t^u \end{pmatrix}$$

$$s_t = H \begin{pmatrix} y_t^n \\ y_{t-1}^n \\ u_t \end{pmatrix} + \begin{pmatrix} \eta_t \\ 0 \end{pmatrix} = \begin{pmatrix} -\kappa & 0 & 0 \\ \theta\rho_y & -\theta\rho_y & 1 \end{pmatrix} \begin{pmatrix} y_t^n \\ y_{t-1}^n \\ u_t \end{pmatrix} + \begin{pmatrix} \eta_t \\ 0 \end{pmatrix}$$

where s_t are the two signals (or "sums") described above.

The Kalman filter yields a recursive formula for expectations $\tilde{E}_t$

$$\tilde{E}_t \begin{pmatrix} y_t^n \\ y_{t-1}^n \\ u_t \end{pmatrix} = F\tilde{E}_{t-1} \begin{pmatrix} y_{t-1}^n \\ y_{t-2}^n \\ u_{t-1} \end{pmatrix} + K_t \left(s_t - HF\tilde{E}_{t-1} \begin{pmatrix} y_{t-1}^n \\ y_{t-2}^n \\ u_{t-1} \end{pmatrix} \right). \tag{2.6}$$

We compute the Kalman-gain matrix K_t numerically, assuming, as is standard in the literature, that the agents' learning problem has already converged such that matrix $K_t = K$ is time-invariant (e.g., Lorenzoni, 2009).

2.3.2 Estimation

We estimate our model on the basis of an indirect inference approach (Gourieroux et al., 1993; Smith, 2008). Indirect inference estimation relies on finding parameters such that an implied moment of the model matches the same moment that characterizes the data—in our case, we match the impulse response functions following a monetary shock displayed in Figure 2.1 above. In the language of Nakamura and Steinsson (2018b) we seek to match an "identified moment" rather than an unconditional moment.

This approach comes with several advantages. First, as explained in Nakamura and Steinsson (2018b), it is relatively robust to misspecification in the structural model because the matching procedure relies only on the part of the model that is needed to generate the particular moment. Second, this approach is also robust to misspecification in the *empirical* model that is used to generate the moment (the "auxiliary model"), because for indirect inference to work, the "auxiliary model need not be correctly specified" (Smith, 2008). For the purpose at hand this matters, because the empirical impulse response functions in Figure 2.1 have been obtained on the basis of identification assumptions which are not generally satisfied by our structural model (e.g., Mertens and Ravn, 2011).[9] Finally, indirect inference is identical to maximum likelihood if the auxiliary model is correctly specified.

We fix the behavioral parameters to reasonable values and only estimate the shock parameters, which we summarize in vector $\varphi = [\rho_u, \rho_y, \sigma_y, \sigma_\eta]'$. These are of particular interest, because the relative size of the variances of the shocks determines the extent of information frictions in our model. Notice that the standard deviation of monetary innovations σ_u is not included in the vector of parameters to be estimated. This is because the Kalman filter output (2.6) only depends on variance (signal-to-noise) *ratios*, but not on the individual levels of the variances. Therefore, without loss of generality we may fix one of the variances in the estimation. We set $\sigma_u = 0.1$.

We chose the following remaining parameter values. We set $\beta = 0.99$, as we assume that a period in the model corresponds to one quarter. Hence, the real interest rate in steady state amounts to one percent per quarter. We use the conventional value for the Taylor coefficient and set $\phi = 1.5$. For the Phillips curve we use a slope of $\kappa = 0.01$, in line with estimates by Gali and Gertler (1999). Finally, for the IES we assume $\theta^{-1} = 0.25$. According to Hall (1988) there is no strong evidence for the IES to be different from zero. Other studies have found higher values (e.g., Smets and Wouters, 2007, report a value of about 0.7).

For the estimation we proceed as follows. For each parameter draw φ we solve the model numerically and simulate a sequence of 234 observations for the (annualized) nominal interest rate, output, the CPI and the nominal exchange rate.[10] We drop the first 100 observations as burn-in period and treat the remaining observations in the same way as the actual time-series data: we run local projections and estimate the impulse response functions for all variables to a monetary policy innovation ε_t^u. Importantly, at this stage we use the same specification for the local projections as in Figure 2.1, that is, a lag structure of $J = 4$ and $K = 4$. We also impose that output and the price level do not respond instantaneously to the shock. We repeat the regression stage 500 times and take the average

[9] Output and prices are not predetermined in our structural model, but following Coibion et al. (2017) we assume this to be the case in our empirical model (2.1). This also rules out direct impulse response matching as, for instance, in Christiano et al. (2005).

[10] In our model, p_t is the producer price index (PPI) whereas the estimation uses on the consumer price index (CPI), which in our model is $cpi_t = (1-\omega)p_t + \omega e_t$ (see the appendix for details). This implies that one additional behavioral parameter, the degree of openness $\omega \in (0, 1)$, needs to be fixed in the estimation. We use $\omega = 0.15$, because imports account for roughly 15% of GDP in the US in our sample.

Parameter	ρ_y	ρ_u	σ_y	σ_η
Estimate	.749(.081)	.960(.003)	.063(.011)	.007(.002)

Statistic	$K_1^{\text{est}}/((\theta\rho_y)/\kappa)$	$K_2^{\text{est}}/1$
Value	.022	.338

Table 2.1: Parameter estimates (standard errors in parentheses) based on indirect inference procedure and implied noise statistics (see the main text for details).

of the impulse responses to eliminate sample noise. Finally, we compute the (weighted) squared distance of the implied impulse responses to the empirical impulse responses from Figure 2.1

$$\hat{\varphi} = argmin_\varphi \; (\hat{\Lambda}^{emp} - \hat{\Lambda}^{sim}(\varphi))'\Sigma^{-1}(\hat{\Lambda}^{emp} - \hat{\Lambda}^{sim}(\varphi)), \tag{2.7}$$

where $\hat{\Lambda}^{emp}$ are the (vectorized) empirical impulse responses, $\hat{\Lambda}^{sim}$ are the simulated impulse responses which depend on the parameter draw φ, and $\hat{\varphi}$ is our estimated vector of parameters. The matrix Σ is a diagonal weighting matrix which contains the variance (point-wise) of the empirical impulse response functions. Therefore, our estimator ensures that the model-implied impulse response functions are as close to the empirical ones as possible, in terms of point-wise standard deviations.

Table 2.1 shows the results. For the natural rate process we estimate an autocorrelation of $\rho_y = 0.749$. In turn, the standard deviation of the innovations is estimated at $\sigma_y = 0.063$. As for the monetary policy shock, we estimate a high autocorrelation of $\rho_u = 0.96$.[11] Lastly, the standard deviation of the noise η_t is estimated to be $\sigma_\eta = 0.007$. We also report standard deviations of our parameter estimates in parentheses.[12]

Figure 2.5 shows the impulse response functions of the estimated structural model (red

[11] In our model, we have abstracted from interest rate smoothing in the Taylor rule. Therefore, the persistence of the federal funds rate observed empirically is absorbed by a high autocorrelation of the monetary shocks. In this sense, our estimates are compatible with earlier estimates by Smets and Wouters (2007).

[12] To compute the standard deviations, we follow Hall et al. (2012) and use

$$\Sigma_\varphi = \Lambda_\varphi \left[\frac{\partial\hat{\Lambda}^{sim}(\varphi)}{\partial\varphi}|_{\varphi=\hat{\varphi}}\right]' \Sigma^{-1}\Sigma_S\Sigma^{-1}\left[\frac{\partial\hat{\Lambda}^{sim}(\varphi)}{\partial\varphi}|_{\varphi=\hat{\varphi}}\right]\Lambda_\varphi,$$

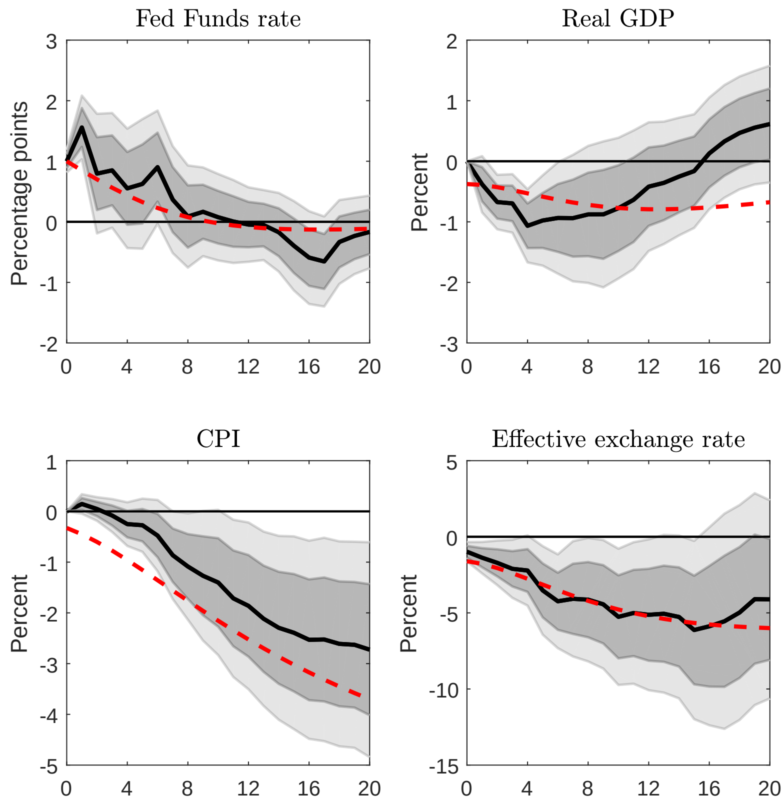

Figure 2.5: Theoretical (model-based) versus empirical impulse response functions (see Figure 2.1). The theoretical impulse response functions are dashed in red. The shocks underlying the theoretical impulse response functions are estimated by using an indirect inference approach, see the description in the text.

dashed lines) and compares them to our baseline empirical impulse response functions from Figure 2.1. We find that the model is able to replicate the empirical patterns well. For example, it is able to generate a hump-shaped response of real GDP. The fact that noisy information models are able to generate humps has already been stressed by Mackowiak and Wiederholt (2015). Importantly, the model response tracks the response for the nominal

where Λ_φ and Σ_S are defined as

$$\Lambda_\varphi \equiv \left(\left[\frac{\partial \hat{\Lambda}^{sim}(\varphi)}{\partial \varphi} |_{\varphi=\hat{\varphi}} \right]' \Sigma^{-1} \left[\frac{\partial \hat{\Lambda}^{sim}(\varphi)}{\partial \varphi} |_{\varphi=\hat{\varphi}} \right] \right)^{-1}, \quad \Sigma_S \equiv \Sigma + \frac{1}{500^2} \sum_{j=1}^{500} \Sigma_j,$$

where Σ_j is the counterpart of matrix Σ in the jth replication of our model-implied impulse response functions. See also Mertens and Ravn (2011) who perform an identical procedure.

exchange rate very well—our key variable of interest. Hence, the estimated model is able to generate *undershooting*, the feature of exchange rate dynamics conditional on monetary policy shocks that stands out in the data.

2.3.3 Measuring the extent of information frictions

We inspect the mechanism by which information frictions impact exchange rate dynamics in the next section. Before doing so, we quantify the extent of information frictions in the model that is implied by our estimates. This assessment reveals to which extent these frictions are *required* in order to generate impulse response functions which we see in the data.

To set the stage, we note that when there is full information, agents perfectly observe the realization of any random variable x_t in the model: $\tilde{E}_t x_t = x_t$. Full information is nested in the model for $\sigma_\eta^2 = 0$ in which case there is no noise in the Phillips curve (2.1): by observing the two signals, private agents can perfectly distinguish changes in potential output y_t^n and monetary policy shocks u_t. Full information implies for the monetary shock u_t, the last row in the Kalman filter output (2.6):

$$u_t = \rho u_{t-1} + \begin{pmatrix} K_1^{\text{full}} & K_2^{\text{full}} \end{pmatrix} \begin{pmatrix} -\kappa & 0 & 0 \\ \theta\rho_y & -\theta\rho_y & 1 \end{pmatrix} \begin{pmatrix} y_t^n - (1+\rho_y) y_{t-1}^n + \rho_y y_{t-2}^n \\ 0 \\ u_t - \rho_u u_{t-1}, \end{pmatrix}.$$

Here $(K_1^{\text{full}}, K_2^{\text{full}})$ denotes the last row of the Kalman matrix K under full information. For this equation to hold, it must be that $K_1^{\text{full}} = (\theta\rho_y)/\kappa$ and $K_2^{\text{full}} = 1$. At the opposite end, when there is *zero* information about the monetary policy shock, it is clear that $K_1^{\text{zero}} = K_2^{\text{zero}} = 0$ for in this case, agents attach zero weight to new information contained in any of the two signals.[13]

This implies that the estimated Kalman filter coefficients lie in the two intervals $K_1^{\text{est}} \in [0, (\theta\rho_y)/\kappa]$ and $K_2^{\text{est}} \in [0, 1]$. It also provides a first interesting statistic about the degree of noisy information implied by our estimates. If K^{est} is estimated to be closer to the upper bound in its interval, there is a small degree of noisy information. Conversely, for

[13]To generate zero information in this model, it is not sufficient to set the noise variance to infinity $\sigma_\eta^2 = \infty$. This is because while in this case, the signal which stems from the Phillips curve becomes uninformative (recall equation (2.1)), agents can still infer about the state of the economy from the signal coming from the Taylor rule (equation (2.4)). Therefore, to have zero inference for the agents about the monetary shock u_t, it must also be that the variance of the natural rate shock is large. Put differently, $\sigma_y^2 = \infty$.

estimates K^{est} closer to zero the degree of noisy information is high.

The last two rows of Table 2.1 show the results. We express K^{est} relative to full information. By using this normalization, numbers closer to one provide a *relative* distance to the case of full information. We obtain 0.022 for the first and 0.338 for the second signal, respectively. Recalling that the first signal comes from the Phillips curve whereas the second signal comes from the Taylor rule, we conclude that agents can infer close to nothing about the monetary policy shock from the Phillips curve, and use about one third of the signal coming from the Taylor rule to update their belief about monetary policy shocks—both indicating a high estimated degree of noisy information. Although based on an entirely different approach and data set, our estimates are thus consistent with the finding of Nakamura and Steinsson (2018a): they find that about two thirds of monetary innovations in the Taylor rule are perceived to be natural-rate innovations, and only one third representing monetary policy shocks.

We may also compute a composite statistic which merges the two previous statistics into one. As described in Coibion and Gorodnichenko (2012), in the presence of two signals, a composite statistic can be obtained by multiplying the Kalman matrix K with the observation matrix H. Because we are interested in the noise when observing monetary shocks, we again evaluate the last row of the resulting matrix (compare equation (2.6)). Our statistic is the last entry in this resulting vector, for this entry determines the weight given in the two signals to monetary policy shocks.

As one can verify, in our case this statistic equals $K_2^{\text{est}} \in [0, 1]$, and is thus the same as the weight given to signals in the Taylor rule discussed above. Here we had found that $K_2^{\text{est}} = 0.338$. Therefore, the overall degree of information processing regarding monetary policy shocks is estimated to be about one third. This is in line with estimates for the noisy information models in Coibion and Gorodnichenko (2012).

2.4 Inspecting the mechanism

In this section we zoom into the details of the transmission mechanism of our model. This allows us to explore how information frictions shape exchange rate dynamics in response to monetary policy shocks. To set the stage, we first consider the full information benchmark. We consider the case of noisy information afterwards. In a last section we dissect the driving forces which shape the estimated impulse response functions in Figure 2.5.

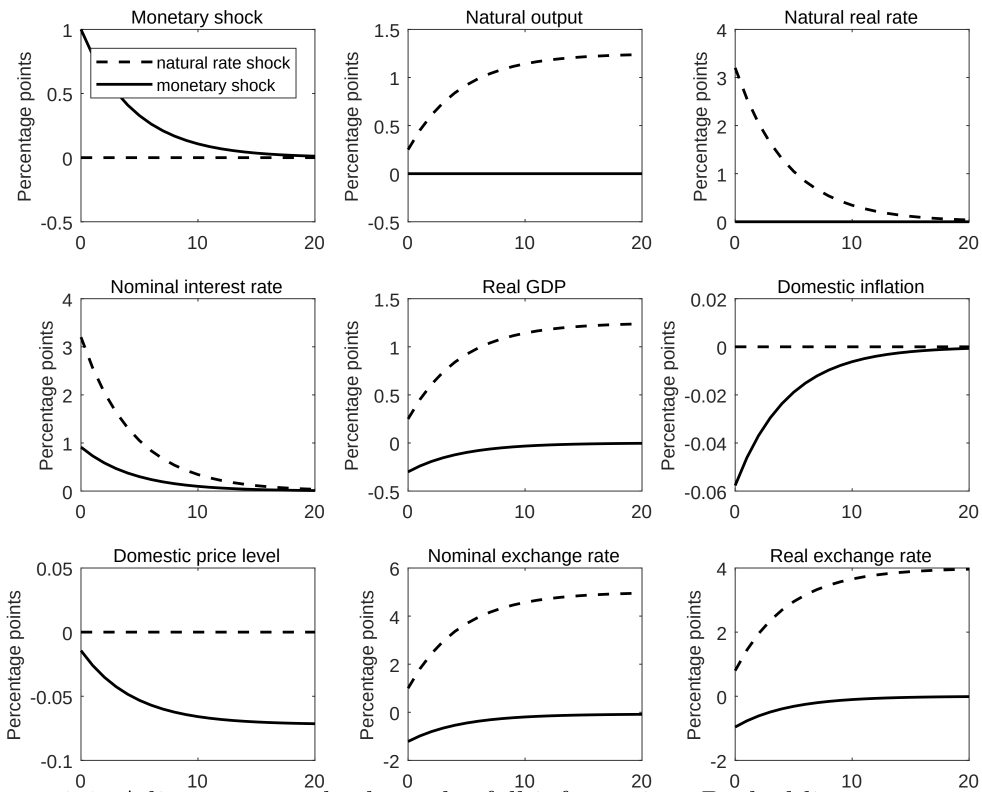

Figure 2.6: Adjustment to shocks under full information. Dashed line: response to natural rate shock; solid line: monetary policy shock. All variables are expressed in percent relative to steady state. Inflation and interest rates are annualized. Parameter values as set/estimated in Section 2.3, except for $\sigma_\eta^2 = 0$ and $\rho_y = \rho_u = 0.8$.

2.4.1 Full information responses

As explained earlier above, our model nests the case of full information for $\sigma_\eta^2 = 0$. For this case, we consider how the economy reacts to a shock to potential output and to a monetary policy shock in turn. Results are shown in Figure 2.6. In order to solve the model numerically, we use the estimated parameters from Section 2.3 except that we set $\sigma_\eta^2 = 0$ as explained before, and that we use $\rho_y = \rho_u = 0.8$ to facilitate the visual comparison of impulse responses across the two shocks. The figure therefore provides a qualitative (not so much a quantitative) illustration of how the model works under full information—a prerequisite for understanding the model with noisy information below.

Focus first on the shock to potential output (dashed), which rises initially until it settles on a permanently higher level. The natural rate rises, foreshadowing the growth path of potential output. In response to the increase of the natural rate, the central bank raises its policy rate. As a result, output follows potential output and inflationary pressure does not arise. This is a standard "divine coincidence" result and underlines the importance of the central bank tracking the natural rate in our model. In doing so, it closes the output gap and stabilizes domestic inflation (see, for instance, Galí, 2015). Observe also that the real exchange rate depreciates permanently in response to the shock, brought about by a permanent nominal depreciation. This is a supply effect: as the supply of domestic goods rises permanently, their price declines on world markets—that is, the real exchange rate depreciates.

Next, we focus on the effect of a monetary policy shock, represented by the solid lines. The central bank tightens interest rates for reasons exogenous to the economy. Potential output and the natural rate are unchanged. Output declines. As the output gap becomes negative, inflation is lowered and, as a result, the price level declines permanently. All of these effects are well known. What is more interesting is the response of the nominal and the real exchange rate. In particular, we note that the nominal exchange rate appreciates in the long term, and that it appreciates by *more* in the short term (see also Figure 2.7 below, which zooms into the response of the nominal exchange rate). Therefore, under full information, our model features exchange rate overshooting, just like in Dornbusch (1976).

Two equations, in particular, govern the nominal exchange rate response. The first is equation (2.2), repeated here for convenience:

$$\theta y_t = e_t + p^* - p_t.$$

This equation determines how the nominal exchange rate reacts in the *long run.* A monetary tightening cannot have an effect on output in the long term. However, because it generates a temporary decline in inflation, the price level p_t declines permanently to a lower level, $p_\infty < p_{-1}$ (recall Figure 2.6). This, in turn, implies that the nominal exchange rate must *appreciate in the long run*, even though the monetary contraction is purely transitory, $e_\infty < e_{-1}$.[14]

The second equation is the UIP condition (2.3), also repeated here for convenience for

[14] Of course, the precise levels of p_∞ and e_∞ are equilibrium objects, determined by the responses of inflation and the nominal exchange rate in the short term.

the full information case:

$$i_t - i^* = E_t \Delta e_{t+1}.$$

Note that, in contrast to equation (2.3), here we use the expectation operator E_t.

A monetary contraction implies a surprise increase of the policy rate at time 0, $i_0 > i^*$, from the Taylor rule (2.4). While by definition, this is unanticipated in period 0, after period 0 all fundamental uncertainty is resolved in the experiment under consideration. This implies that $E_t \Delta e_{t+1} = \Delta e_{t+1}$ for all $t \geq 0$, because under full information, agents are not making an expectational error. Dornbusch (1976)'s overshooting result follows immediately: $i_t - i^* > 0$ implies that $\Delta e_{t+1} > 0$, that is, the nominal exchange rate must *depreciate over time*. Because the exchange rate appreciates in the long run, $e_\infty < e_{-1}$, the way in which both are compatible is that in the initial period 0, the exchange rate *overshoots*, $e_0 < e_\infty < e_{-1}$.

2.4.2 Exchange rate dynamics when information is noisy

Now, we turn to the case of information frictions. Once the monetary policy shock hits, private agents observe a policy rate rise. However, they are unable to tell whether this represents a rise in the natural rate to which monetary policy responds, or a monetary policy shock. They are therefore unsure whether in the long term, the nominal exchange rate is going to *depreciate* or to *appreciate* (recall Figure 2.6). Because the nominal exchange rate is a forward looking variable, its current response reflects this uncertainty. For example, if agents attach a high probability to the rate rise reflecting a change in the natural rate, the nominal exchange rate may initially depreciate. As agents realize that the policy rate rise represents a monetary contraction, the exchange rate starts to appreciate over time. Generally, the model with information frictions can thus account for overshooting, undershooting or delayed overshooting—depending on the model parameters which determine how market participants process information.

This intuition can be made precise formally. We repeat equation (2.3) for convenience

$$i_t - i^* = \tilde{E}_t \Delta e_{t+1}.$$

Unlike under full information, agents in noisy information models make expectational errors even absent any fundamental surprises, that is, even *after* the shock has hit in period zero.

In fact, in noisy information models, the expectational error only *converges* to zero in the long term (Coibion and Gorodnichenko, 2012). Formally, following a shock to u_t in the initial period, $\Delta e_{t+1} \neq \tilde{E}_t \Delta e_{t+1}$ for all $t \geq 0$. Letting $v_{t+1} \equiv \Delta e_{t+1} - \tilde{E}_t \Delta e_{t+1}$ denote the expectational error, we may rewrite the last equation as

$$i_t - i^* + v_{t+1} = \Delta e_{t+1}. \tag{2.1}$$

Equation (2.1) illustrates why our model may not predict exchange rate overshooting to the extent that information is noisy. Even when the policy rate i_t rises, a negative enough expectational error,

$$v_{t+1} < 0, \tag{2.2}$$

can imply a nominal *appreciation over time* even though the domestic currency carries a high interest rate. That the expectational error must indeed be negative can again be understood from the exchange rate response in Figure 2.6. Under noisy information, agents initially expect the exchange rate to appreciate by less than under full information, from previous arguments. Over time, as they learn about the monetary policy shock, they realize that the exchange rate will appreciate. This implies that $\tilde{E}_t \Delta e_{t+1} > e_{t+1}$, or that $v_{t+1} < 0$.

We illustrate how the nominal exchange rate response changes once we gradually adjust the noise variance σ_η^2 from zero to a positive value. The result is shown in Figure 2.7, the right panel. When $\sigma_\eta = 0$, the nominal exchange rate response is characterized by overshooting, as in Figure 2.6. As σ_η is raised, the impact response of the exchange rate is weakened and—for some time—the exchange rate *is appreciating* over time rather than depreciating. For a low level of information frictions, the exchange rate response is thus characterized by *delayed overshooting*. Instead, as information frictions become more severe, the exchange rate response changes from delayed overshooting to undershooting.

Finally, the left panel in Figure 2.7 shows that, whatever the noise variance σ_η, the nominal interest rate response is virtually identical. This highlights that, quite independently of the nominal interest rate response, our model can explain varying shapes of the nominal exchange rate response, depending on the degree of information frictions and therefore on the size of the expectational error (see equation (2.1)).

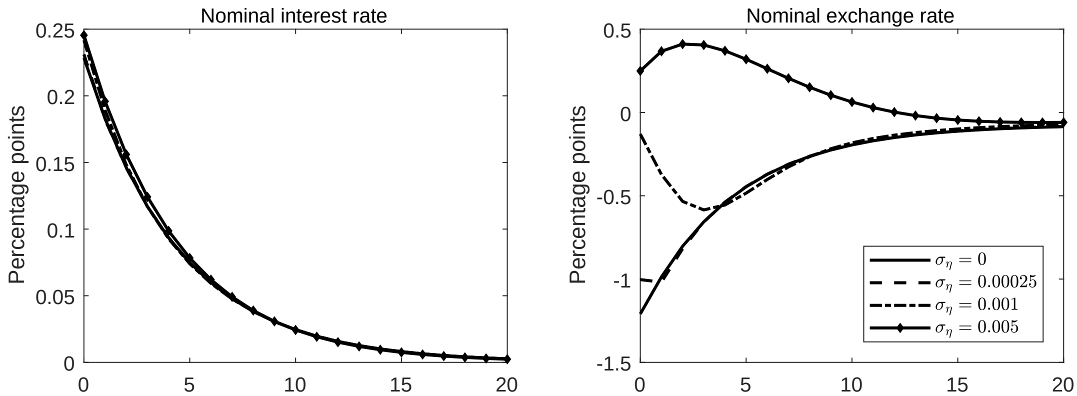

Figure 2.7: Impulse response functions of the nominal interest rate i_t and the nominal exchange rate e_t following a monetary policy shock. Parameters as estimated in Section 2.3, except σ_η^2, which we vary from zero to a positive number, and $\rho_y = \rho_u = 0.8$. Compare Figure 2.6.

2.4.3 Dissecting the estimated nominal exchange rate response

We now dissect the nominal exchange rate response of the estimated model, shown in Figure 2.5, as we identify its underlying drivers. Step by step we uncover how our empirical undershooting result can be explained by information frictions.

The upper-left panel in Figure 2.8 decomposes the nominal exchange rate response according to equation (2.1). Here we therefore split the (change in the) nominal exchange rate (solid lines) into nominal interest differential (diamonds) plus the expectational error (dashed lines).

Under full information, the expectational error would be zero in all periods except in period 0, as argued above. The solid line and the line with diamonds would thus *coincide*: a high interest rate would be accompanied by ongoing nominal depreciation. Not so under noisy information. In this case, the expectational error v_{t+1} is negative (see equation (2.2)), which *drags down* the response of exchange rate changes relative to full information. In our estimated model, the expectational-error effect is strong enough to overturn the *sign* of the change of the exchange rate response from positive to negative for the entire horizon of the response. Rather than depreciating, the nominal exchange rate *appreciates* over time despite a high interest rate. This is the core of our undershooting result.

The remaining three panels show the source of the expectational error v_{t+1} or, informally,

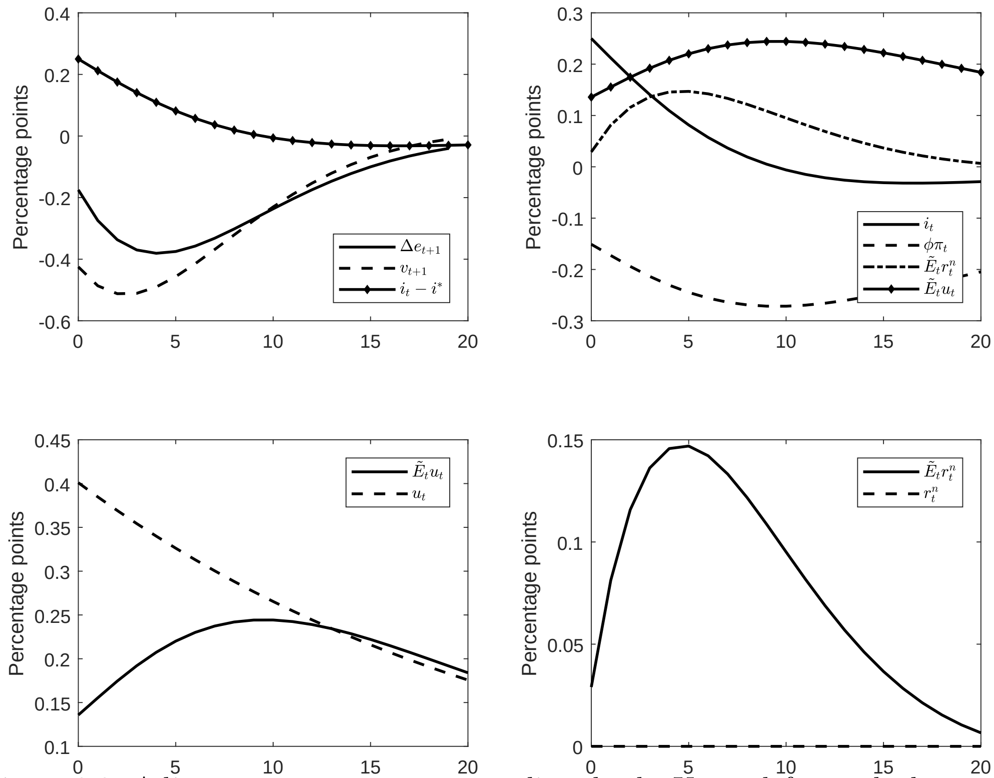

Figure 2.8: Adjustments to a monetary policy shock. Upper-left panel: decomposition by using equation (2.1). Upper-right panel: components of the Taylor rule (2.3). Lower panels: actual versus perceived evolution of monetary policy shock u_t and natural rate r_t^n.

the source of private agents' expectational error. The upper-right panel decomposes the Taylor rule (2.4) into its individual components, after applying the expectations operator:

$$i_t = \tilde{E}_t r_t^n + \phi\pi_t + \tilde{E}_t u_t, \tag{2.3}$$

where we use that $\tilde{E}_t i_t = i_t$ and $\tilde{E}_t \pi_t = \pi_t$, because both i_t and π_t are perfectly observed. As agents observe a policy rate rise i_t, they (partially) mistake this to be a policy response to a natural rate rise even though the natural rate has not changed. In fact, a significant share of the probability weight is initially put on a natural rate disturbance.

The two lower panels reveal this in more detail as they show the perceived response $\tilde{E}_t u_t$ and $\tilde{E}_t r_t^n$ versus the *actual* response for these two variables. By looking at the actual response, we see that the underlying dynamics in Figure 2.5 is a monetary policy shock which initially rises to 0.43 percentage points, which slowly returns to zero due to a high estimated autocorrelation. Instead, the natural rate stays constant at zero. By observing the response of the economy over time, private agents update their beliefs and adjust their estimates of the two shocks accordingly. However, it takes more than five years (twenty quarters) until private agents put their estimate for the natural rate to the true value of zero, and about three years (twelve quarters) until the private agents' perception and the actual evolution of the monetary policy shock roughly coincide.

We conclude that, for the model to match the empirically observed impulse response functions following a monetary contraction, the required degree of information friction on monetary policy shocks is substantial.

2.5 External validation

In the previous section we have shown that a model with information frictions is able to account for the empirical impulse responses to a monetary policy shock. In particular, it is able to account for the extent of exchange rate undershooting that characterizes our identified moments. In our estimation, we determine parameter values so as to minimize the distance between the model predictions and the identified moments. Against this background it is interesting to confront the predictions of the model with additional evidence, notably evidence that is not used in the estimation procedure.

2.5.1 The exchange rate response to supply shocks

In our small open economy model a shock to the natural rate that raises potential output induces the exchange rate to depreciate in the long run. In order to test this prediction of the model we estimate the response of the effective nominal exchange rate of the US dollar to technology shocks. For this purpose, we employ once more our empirical model (2.1) and project the change in the exchange rate at various horizons on TFP innovations as provided by Fernald (2014).[15]

Figure 2.9 shows the result. In the left panel we show the result for a sample that is as close as possible to our baseline. Specifically, it covers the period 1976 - 2007. In this case we find that the exchange rate depreciates in response to TFP shocks, but the effect is not

[15]The shock series represents the change in TFP while accounting for changes in utilization.

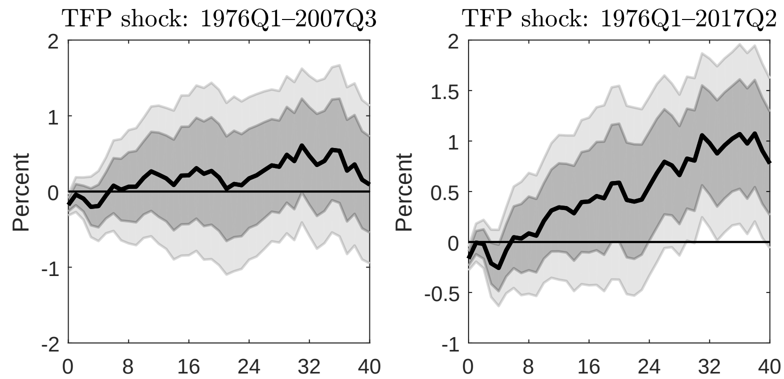

Figure 2.9: Exchange rate response to a TFP shock. Solid line represents point estimate, shaded areas correspond to 68% and 90% confidence intervals, respectively. Time (horizontal axis) is in quarters. TFP shocks series provided by Fernald (2014).

statistically significant. In the right panel, we use a longer sample. In this case, we do find a significant depreciation, in line with the predictions of the model.

2.5.2 Monetary policy and growth: reassessing the information effect

A striking observation by Nakamura and Steinsson (2018a) is that in response to a monetary policy surprise—identified on the basis of high frequency data—survey estimates of expected output growth increase. This observation is pivotal in order to motivate their analysis of the information effect. In our analysis we rely on the measure of monetary policy shocks put forward by Romer and Romer. In this case, shocks are by construction orthogonal to the information set of the federal reserve. As a result these shocks should not convey new information about the current or expected state of the economy to market participants.

Still, it is instructive to assess how growth expectations respond to monetary policy shocks in the context of our analysis. For this purpose we consider quarterly observations from the Survey of Professional Forecasters. The left panel of Figure 2.10 correlates the change in growth expectations one year ahead with Romer-Romer shocks. As in Nakamura and Steinsson (2018a) who consider monthly observations in Blue Chip survey expectations, we find a positive association. A regression yields a significant slope coefficient of 0.24.

Taken at face value, one may conjecture that the Romer-Romer shock is not a genuine monetary policy shock but instead contains some additional information about the economy. We can use the estimated model to assess this conjecture. Specifically, in the right panel

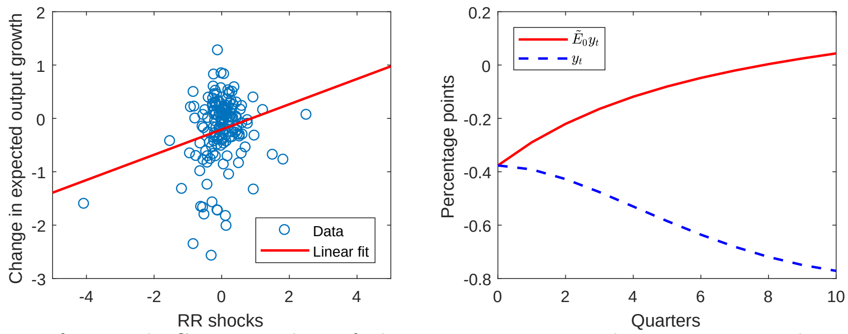

Figure 2.10: Left panel: Scatter plot of changes in expected output growth over next year (quarterly SPF) and Romer, Romer (RR) shocks, slope of linear regression is 0.24. Right panel: Expectations of output at period zero ($\tilde{E}_0 y_t$) and realized output (y_t) in response to a monetary policy shock both in percentage points. The parameters are as estimated in Section 2.3.

of Figure 2.10 we display again the impulse response of output to the monetary policy shock (dashed line). In addition, we also plot the expectation of the future path of output on impact, that is, just after the shock materializes (solid line). We find that market participants expect output to grow over time. This is because they assign a high probability to the possibility that the interest rate increase represents a response to the natural rate, even thought the economy is subjected to a monetary policy shock. Hence, we conclude that rising growth expectations do not necessarily imply that monetary policy surprises carry proper news about the state of the economy. In an economy with information frictions market participants may revise their growth expectations upward simply because they do not know the true nature of the shock.

2.6 Conclusion

A number of recent contributions have highlighted the importance of information frictions in order to account for expectations data and related macroeconomic phenomena. In this paper, we study how information frictions impact exchange rate dynamics. This is a first order issue in light of Dornbusch's overshooting hypothesis where expectations take center stage and, importantly, are assumed to adjust instantaneously to shocks.

And, indeed, we find that the exchange rate undershoots in response to monetary policy shocks if information frictions are pervasive—thereby overturning Dornbusch (1976)'s original result. Specifically, we put forward a small open economy model with information frictions. In our model agents do neither observe monetary policy shocks nor the natural rate directly. Market participants thus attach some probability to the possibility that an increase in the policy rate is an endogenous policy response to the natural rate, rather than a monetary policy shock. An increase in the natural rate signals rising potential output which comes with an exchange rate depreciation. Hence, for as long there is uncertainty about the true nature of the shock the exchange rate response is muted.

We also provide evidence for undershooting as we estimate the effect of US monetary policy shocks on the exchange rate and other variables of interest. Specifically, we use local projections to obtain impulse response functions on which we rely, in turn, to estimate the structural model on the basis of an indirect inference procedure. The degree of information friction implied by the estimated model is economically important, and strictly necessary (in our model) to explain the observed undershooting response. This testifies once more to the importance of information frictions when it comes to accounting for key macroeconomic phenomena.

2.A Economic environment

Here we describe the non-linear model in some detail, and present details on the linearization. Much of the exposition is drawn from Kriwoluzky et al. (2013).

2.A.1 Non-linear model

Final Good Firms The final consumption good, C_t, is a composite of intermediate goods produced by a continuum of monopolistically competitive firms both at home and abroad. We use $j \in [0,1]$ to index intermediate goods. Final good firms operate under perfect competition and purchase domestically produced intermediate goods, $Y_t(j)$, as well as imported intermediate goods, $Y_{I,t}(j)$. Final good firms minimize expenditures subject to the following aggregation technology

$$C_t = \left[(1-\omega)^{\frac{1}{\sigma}} \left(\left[\int_0^1 Y_t(j)^{\frac{\epsilon-1}{\epsilon}} dj \right]^{\frac{\epsilon}{\epsilon-1}} \right)^{\frac{\sigma-1}{\sigma}} + \omega^{\frac{1}{\sigma}} \left(\left[\int_0^1 Y_{I,t}(j)^{\frac{\epsilon-1}{\epsilon}} dj \right]^{\frac{\epsilon}{\epsilon-1}} \right)^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}},$$

$$ \tag{2.A.1} $$

where $\sigma > 0$ is the trade price elasticity. The parameter $\epsilon > 1$ measures the price elasticity across intermediate goods produced within the same country, while $\omega \in (0, 1)$ measures the weight of imports in the production of final consumption goods—a value lower than one corresponds to home bias in consumption.

Expenditure minimization implies the following price indices for domestically produced intermediate goods and imported intermediate goods, respectively,

$$ P_t = \left(\int_0^1 P_t(j)^{1-\epsilon} di \right)^{\frac{1}{1-\epsilon}}, \qquad P_{I,t} = \left(\int_0^1 P_{I,t}(j)^{1-\epsilon} di \right)^{\frac{1}{1-\epsilon}}. \tag{2.A.2} $$

By the same token, the consumption price index is

$$ CPI_t = \left((1-\omega) P_t^{1-\sigma} + \omega P_{I,t}^{1-\sigma} \right)^{\frac{1}{1-\sigma}}. \tag{2.A.3} $$

Regarding the rest of the world (ROW), we assume an isomorphic aggregation technology. Further, the law of one price is assumed to hold at the level of intermediate goods such that

$$ P_{I,t} = P_t^* \mathcal{E}_t, \tag{2.A.4} $$

where $\mathcal{E}_t$ denotes the nominal exchange rate (the price of foreign currency in terms of domestic currency). P_t^* denotes the price index of imports measured in foreign currency. We also define the terms of trade and the real exchange rate as

$$ S_t = \frac{P_{I,t}}{P_t}, \; Q_t = \frac{P_t^* \mathcal{E}_t}{CPI_t}, \tag{2.A.5} $$

respectively. While the law of one price holds throughout, deviations from purchasing power parity are possible in the short run, due to home bias in consumption.

Intermediate Good Firms Intermediate goods are produced on the basis of the following production function: $Y_t(j) = H_t(j)$, where $H_t(j)$ measures the amount of labor employed by firm j. Intermediate good firms operate under imperfect competition. We assume that price setting is constrained exogenously à la Calvo. Each firm has the opportunity to change its price with a given probability $1-\xi$. Given this possibility, a generic firm j will set $P_t(j)$ in order to solve

$$\max \tilde{E}_t \sum_{k=0}^{\infty} \xi^k \rho_{t,t+k} \left[Y^d_{t,t+k}(j) P_t(j) - W_{t+k} H_{t+k}(j) \right], \tag{2.A.6}$$

where $\rho_{t,t+k}$ denotes the nominal stochastic discount factor and $Y^d_{t,t+k}(j)$ denotes demand in period $t+k$, given that prices have been set optimally in period t. Note that expectations have a tilde $\tilde{E}_t$ to indicate the presence of incomplete information.

Households The domestic economy is inhabited by a representative household that ranks sequences of consumption and labor effort as

$$\tilde{E}_t \sum_{k=0}^{\infty} \beta^k \left(\frac{C^{1-\theta}_{t+k}}{1-\theta} - \frac{H^{1+\varphi}_{t+k}}{1+\varphi} \right), \quad \beta \in (0,1) \tag{2.A.7}$$

The household trades a complete set of state-contingent securities with the rest of the world. Letting Ξ_{t+1} denote the payoff in units of domestic currency in period $t+1$ of the portfolio held at the end of period t, the budget constraint of the household is given by

$$W_t H_t + \Upsilon_t - P_t C_t = \tilde{E}_t \rho_{t,t+1} \Xi_{t+1} - \Xi_t, \tag{2.A.8}$$

where Υ_t denote lump-sum profits of intermediate good firms.

Monetary policy Domestic monetary policy is specified by an interest rate feedback rule. Defining the one-period interest rate as $I_t \equiv 1/\tilde{E}_t(\rho_{t,t+1})$, we posit

$$I_t = R^n_t \Pi^{\phi}_t U_t, \quad \phi > 1, \tag{2.A.9}$$

where $\Pi_t = P_t/P_{t-1}$ measures domestic inflation and (here as well as in the following), R^n_t is the natural rate and U_t is a monetary policy shock.

Market clearing At the level of each intermediate good, supply equals demand of final good firms and the ROW:

$$Y_t(j) = Y_t^d(j) = \left(\frac{P_t(j)}{P_t}\right)^{-\epsilon} \left(\frac{P_t}{CPI_t}\right)^{-\sigma} ((1-\omega)C_t + \omega S_t^{\sigma} C_t^*)\,, \tag{2.A.10}$$

where C_t^* denotes consumption in the ROW. It is convenient to define an index for aggregate domestic output:

$$Y_t = \left(\int_0^1 Y_t(j)^{(\epsilon-1)/\epsilon} dj\right)^{\epsilon/(\epsilon-1)} .$$

Substituting for $Y_t(j)$ using (2.A.10) gives the aggregate relationship

$$Y_t = \left(\frac{P_t}{CPI_t}\right)^{-\sigma} ((1-\omega)C_t + \omega S_t^{\sigma} C_t^*)\,. \tag{2.A.11}$$

2.A.2 Equilibrium conditions and the linearized model

In the following, we use lower-case letters to denote the log of a variable. Variables in the ROW are assumed to be constant throughout.

Price indices The terms of trade, the law of one price, the CPI, CPI inflation and the real exchange rate can be written as

$$\begin{aligned} s_t &= p_{I,t} - p_t, & (2.A.12)\\ p_{I,t} &= p^* + e_t, & (2.A.13)\\ cpi_t &= (1-\omega)p_t + \omega p_{I,t} = p_t + \omega s_t, & (2.A.14)\\ \Delta cpi_t &= \pi_t + \omega \Delta s_t, & (2.A.15)\\ q_t &= (1-\omega)s_t. & (2.A.16) \end{aligned}$$

Intermediate good firms The demand for intermediate good (j) is given by

$$Y_t(j) = \left(\frac{P_t(j)}{P_t}\right)^{-\varepsilon} Y_t, \tag{2.A.17}$$

so that

$$\int_0^1 Y_t(j)dj = \zeta_t Y_t, \tag{2.A.18}$$

where $\zeta_t = \int_0^1 \left(\frac{P_t(j)}{P_t}\right)^{-\varepsilon} dj$ measures price dispersion. Aggregation gives

$$\zeta_t Y_t = \int_0^1 H_t(j)dj = H_t. \tag{2.A.19}$$

A first order approximation is given by $y_t = h_t$.

The first order condition to the price setting problem is given by

$$\tilde{E}_t \sum_{k=0}^{\infty} \xi^k \rho_{t,t+k} \left[Y^d_{t,t+k}(j) P_t(j) - \frac{\varepsilon}{\varepsilon - 1} W_{t+k} H_{t+k} \right] = 0. \tag{2.A.20}$$

Linearizing (2.A.20) around zero inflation, one obtains a variant of the New Keynesian Phillips curve (see, e.g., Galí and Monacelli, 2005a):

$$\pi_t = \beta \tilde{E}_t \pi_{t+1} + \lambda \hat{mc}^r_t, \tag{2.A.21}$$

where $\lambda := (1 - \xi)(1 - \beta\xi)/\xi$ and marginal costs are defined in real terms, deflated with the domestic price index

$$mc^r_t = w_t - p_t = w^r_t + \omega s_t. \tag{2.A.22}$$

Here, a hat $\hat{}$ indicates log-deviation from steady state, and $w^r_t = w_t - cpi_t$ is the real wage deflated with the CPI.

Households The first order conditions in deviations from steady state are

$$w^r_t = w_t - cpi_t \quad = \quad \theta c_t + \varphi h_t, \tag{2.A.23}$$

$$c_t \quad = \quad \tilde{E}_t c_{t+1} - \frac{1}{\theta}(i_t - \tilde{E}_t \Delta cpi_{t+1} - \rho), \tag{2.A.24}$$

where $\rho = -\log(\beta) > 0$. Risk sharing implies that consumption is tightly linked to the real exchange rate (see, e.g., Galí and Monacelli, 2005a)

$$\theta c_t = q_t. \tag{2.A.25}$$

Monetary policy Rewriting the interest rate feedback rule gives

$$i_t = r_t^n + \phi \pi_t + u_t. \tag{2.A.26}$$

Equilibrium Linearizing the good market clearing condition (2.A.11) yields

$$y_t = (2-\omega)\sigma\omega s_t + (1-\omega)c_t, \tag{2.A.27}$$

where we use (2.A.12)-(2.A.15) and set $c^* = 0$.

Key equations We show how to obtain equations (2.1)-(2.3) from the main text (the New Keynesian Phillips curve, the risk sharing condition and the UIP condition).

Combine good market clearing (2.A.27), risk sharing (2.A.25) and the definition of the real exchange rate (2.A.16) to obtain

$$y_t = \frac{1}{\theta} \underbrace{(1+\omega(2-\omega)(\sigma\theta-1))}_{:=\varpi} s_t, \tag{2.A.28}$$

We assume that $\sigma = 1/\theta$ (the so called Cole-Obstfeld condition), in which case $\varpi = 1$. Rearrange to obtain

$$s_t = \theta y_t. \tag{2.A.29}$$

Combine with equations (2.A.12) and (2.A.13) to obtain equation (2.2) in the main text.

Rewrite the Euler equation (2.A.24)

$$\begin{aligned} c_t &= \tilde{E}_t c_{t+1} - \frac{1}{\theta}(i_t - \tilde{E}_t(\pi_{t+1} + \omega\Delta s_{t+1}) - \rho) && (2.A.30) \\ &= \tilde{E}_t c_{t+1} - \frac{1}{\theta}(i_t - \tilde{E}_t\pi_{t+1} + \omega\theta\tilde{E}_t\Delta y_{t+1} - \rho), && (2.A.31) \end{aligned}$$

where we use (2.A.15) in the first line and (2.A.29) in the second.

Combine (2.A.29) with (2.A.25) and (2.A.16) to obtain

$$c_t = (1 - \omega) y_t. \tag{2.A.32}$$

Use this expression to substitute for consumption in (2.A.31)

$$y_t = \tilde{E}_t y_{t+1} - \frac{1}{\theta}(i_t - \tilde{E}_t \pi_{t+1} - \rho), \tag{2.A.33}$$

which is the dynamic IS curve. The same equation holds in ROW. Using that p^* is constant and therefore that $\pi^* = 0$, and using that $y^* = c^* = 0$, this implies $i^* = \rho$. Using this and combining the DIS curve (2.A.33) with equation (2.2) from the main text, we obtain the UIP condition (2.3) from the main text.

Finally, we rewrite the Phillips curve (2.A.21). We use (2.A.23), (2.A.29), (2.A.32) and production technology $y_t = h_t$ to rewrite marginal cost

$$mc_t^r = w_t^r + \omega s_t = \theta c_t + \varphi h_t + \omega s_t = (\theta + \varphi) y_t. \tag{2.A.34}$$

Insert this into the Phillips curve (2.A.21) and define $\kappa \equiv \lambda(\theta + \varphi)$ to obtain equation (2.1) in the main text.

Chapter 3

Is the German current account surplus too large?

Joint with Marc Faupel

3.1 Introduction

The German current account in percent of GDP has remained at an elevated level since 2007 with a peak of 10% in 2016 as illustrated in Figure 3.1. Given such high numbers there is an ongoing debate on whether the German current account surplus is too large. For example, The Economist (2017) argues that the large surplus amplified the crises in some southern-European countries and that it unreasonably strains the global trading system. Further, IMF (2017) sees the surplus above what is implied by economic fundamentals and recommends policies to narrow the surplus such as increasing public investment. The German Federal Ministry of Finance (2017) on the contrary, regards the surplus as a market outcome which is driven, for example, by increased savings of an aging population.

We use the so called intertemporal approach to the current account as a benchmark to judge whether the German current account surplus is too large. According to this approach, the current account predicts changes in the present value of future income (net of investment and government spending). The underlying mechanism of this prediction is the desire to smooth consumption of a representative household. Hence, a country runs a current account surplus when it expects future income to decline. The cross-equation restriction of the intertemporal model on a vector autoregression (VAR) has been tested empirically several times (for early test see e.g. Sheffrin and Woo, 1990; Otto, 1992). It turns out that the predicted current account often resembles the true current account strikingly well, although the cross-equation restriction is frequently rejected by the data.[1] More recently, Bergin and Sheffrin (2000) improve the fit of the model by extending the simple representative

[1] See for example the discussion in Obstfeld and Rogoff (1996). For a comprehensive survey of papers that tested the intertemporal approach see Herzberg (2015).

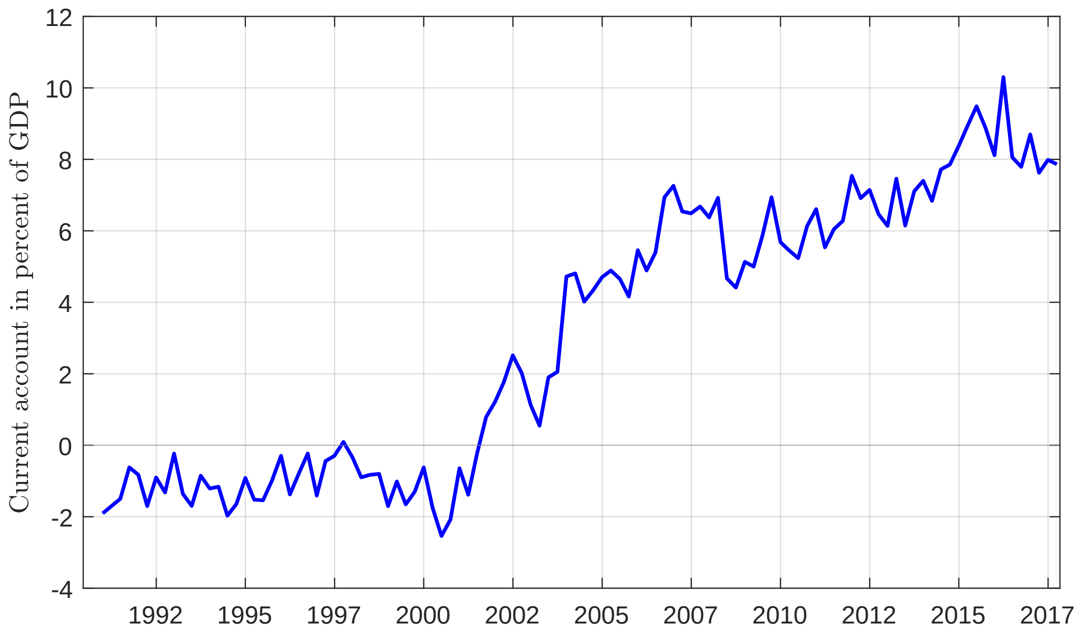

Figure 3.1: German current account in percent of GDP, seasonally adjusted, range is from 1991-Q1 to 2017-Q4; source: OECD.

household model to include non-separable utility between tradable and non-tradable goods and a time-varying world interest and exchange rate.

In order to answer our question, we use the intertemporal approach of Bergin and Sheffrin (2000). For this purpose we estimate a VAR for Germany and find that the cross-equation restrictions of the intertemporal model cannot be rejected. Hence, it cannot be rejected that the German current account is explained by the intertemporal approach. Further, the current account which is predicted by the cross-equation restrictions matches the large surplus. Therefore, the German current account surplus appears justified from the perspective of the intertemporal approach.

We are not the first ones to analyze the large German current account surplus. Kollmann et al. (2015) estimate a large-scale DSGE model to quantify which of the about 40 shocks drive the German current account. They find that shocks to the German savings rate account for about half of the surplus after 2008. In his comment to Kollmann et al. (2015), Müller

(2015) discusses potential driving forces of the savings shock including the consumption smoothing incentive within the intertemporal model. For this purpose he analyzes to which extend a further prediction of the intertemporal model explains the current account surplus. The prediction is that a country runs a current account surplus if it expects a declining share in world output (Engel and Rogers, 2006). He finds that the surplus can be justified by this prediction since the implied world output share of the intertemporal model for Germany is in line with its projected path by the OECD.

Related to our analysis Campa and Gavilan (2011) evaluate the sustainability of the external imbalances of some countries of the European Monetary Union using the setup put forward in Bergin and Sheffrin (2000). To this end they compute whether growth expectations that drive the current account are in line with historical growth rates. For example, Spain's growth expectations appear to have been overoptimistic prior to the financial crisis.[2] Bussiere et al. (2018) also use the model of Bergin and Sheffrin (2000) to analyze the current account of some major advanced economies (G7 minus the US). They, however, do not test the intertemporal model as we do here but focus on the sources of current account fluctuations.

The remainder of the paper is structured as follows. Section 3.2 presents the literature on the intertemporal approach in more detail, briefly discusses critique put forward against the intertemporal approach and motivates our choice to use the model of Bergin and Sheffrin (2000). Section 3.3 shortly describes their model. Section 3.4 presents the data and discusses the parameterization. We discuss our results in Section 3.5. In Section 3.6 we conduct a Monte Carlo analysis to shed further light on our results. Finally, Section 3.7 concludes.

3.2 Related literature

The intertemporal approach to the current account dates back to Sachs (1981, 1982) who highlighted that the current account should not be analyzed in a static environment. Forward looking consumers will take expectations of their future income into account and adjust their consumption/savings already today. This prediction is analogue to the permanent income hypothesis. In consequence, methods developed to test the permanent income hypothesis (Campbell, 1987; Campbell and Shiller, 1987) have then been applied within the intertemporal approach to the current account. As already mentioned in the

[2] The authors do not conduct this exercise for Germany since they reject the intertemporal model for Germany. This might be due to the shorter sample in their analysis which runs from 1991Q1-2005Q4.

introduction, early tests found the cross-equations restrictions of the intertemporal model frequently rejected although the predicted current account had a good visual fit to the actual current account. These early tests were conducted in a "simple model" of a small open economy in which households receive an exogenous stochastic endowment and can borrow an internationally traded risk-free bond at a constant world interest rate.

The simple intertemporal model has been extended in many directions. For example, Ahmed (1986), Ahmed and Rogers (1995) and Bussière et al. (2010) among others study the role of government spending shocks on the current account. Bergin and Sheffrin (2000) take a different approach and analyze a version of the intertemporal model with tradable and non-tradable goods together with a time-varying world interest and exchange rate. Schmitt-Grohé and Uribe (2003) compare a variety of extensions that make the intertemporal model stationary. For instance, they consider a debt-elastic interest-rate premium within the intertemporal framework. In order to understand why the simple intertemporal model is frequently rejected, Nason and Rogers (2006) conduct a Monte Carlo analysis with the above extensions for Canadian data. They confirm that extending the model with stochastic interest rates—as in Bergin and Sheffrin (2000)—helps to explain the rejection of the simple intertemporal model for Canada. Further, they find that a combination of an internalized debt-elastic interest rate premium together with a stochastic world interest rate matches the Canadian data best.

The intertemporal approach abstracts from valuation effects of the net foreign asset position. Gourinchas and Rey (2014) consider this neglect as responsible for the rejection of the intertemporal model. To illustrate this, the change in the net foreign asset position can be decomposed as follows:[3]

$$NFA_t - NFA_{t-1} = CA_t + VA_t, \tag{3.2.1}$$

where NFA_t denotes the net foreign asset position at market values of a given country at the end of period t. The current account is given by CA_t while VA_t denotes the valuation adjustment. The intertemporal model requires that a country with a current account deficit has to run future current account surpluses. Instead, when valuation effects are considered, the adjustment in the net foreign asset position can also be brought about by a revaluation of net foreign assets (Gourinchas and Rey, 2007). For example, if the liabilities

[3]This representation still abstracts from changes in the capital account, unilateral transfers and the statistical discrepancy, all of which are typically small for industrialized countries, see Gourinchas and Rey (2014).

of a given country are mainly denoted in the domestic currency and the foreign assets are mainly denoted in foreign currency, a depreciation of the exchange rate reduces the need to run future current account surpluses. Of course, the adjustment can also happen by the revaluation of the market value of certain assets. Such adjustments are absent in Bergin and Sheffrin (2000) whose model contains only a single asset. However, the return of this asset may change stochastically and in this respect approximates valuation effects. Further, IMF (2018) estimates average valuation changes in the German international investment position for the period of 2012-2016 to be about only 1% of GDP.[4] We consider this as tolerable.

Since the focus of this paper is to which extent the German current account surplus can be explained by the intertemporal approach, we consider the model by Bergin and Sheffrin (2000) best suited for our analysis. First, their extensions improved the fit of the intertemporal model. Further, the inclusion of a time-varying world interest and exchange rate approximates valuation effects.

3.3 The intertemporal model of Bergin and Sheffrin

In this section we briefly present the model of Bergin and Sheffrin (2000). In their model of a small open economy a representative household maximizes expected lifetime utility by choosing consumption and debt. Formally, the optimization problem is given by:

$$\max E_0 \sum_{t=0}^{\infty} \beta^t U(C_{Tt}, C_{Nt})$$

$$\text{s.t.} \quad Y_t - (C_{Tt} + P_t C_{Nt}) - I_t - G_t + r_t B_{t-1} = B_t - B_{t-1}, \text{ with } B_0 \text{ given}, \tag{3.3.1}$$

where C_{Tt} denotes consumption of traded and C_{Nt} consumption of non-traded goods. Y_t is endowment, I_t is Investment, G_t is government consumption. These variables are measured in terms of traded goods. P_t denotes the relative price of non-traded goods in terms of traded goods. Further, the stock of external assets is denoted by B_t, while r_t is the net world real interest rate in terms of traded goods and may vary exogenously. For the subjective discount factor it holds that $\beta \in (0, 1)$. Ponzi schemes are not permitted. The utility

[4]See Figure 3 in their technical supplement III.B.

function is of the constant relative risk aversion type:

$$U(C_{Tt}, C_{Nt}) = \frac{1}{1-\sigma} \left(C_{Tt}^{a} C_{Nt}^{1-a} \right)^{1-\sigma}, \tag{3.3.2}$$

where $\sigma > 0$ and $0 < a < 1$. Further, total consumption expenditure may be written as $C_t = C_{Tt} + P_t C_{Nt}$. Optimal policy implies the following variant of the consumption-Euler equation:[5]

$$E_t \Delta c_{t+1} = \gamma E_t r_{t+1}^{*}, \tag{3.3.3}$$

where $c_t = \log C_t$ and $\Delta c_{t+1} = c_{t+1} - c_t$. Parameter γ denotes the intertemporal elasticity of substitution with $\gamma = 1/\sigma$ while r^* is given by:

$$r_t^{*} = r_t + \left[\frac{1-\gamma}{\gamma} (1-a) \right] \Delta p_t + \text{constant}, \tag{3.3.4}$$

with $p_t = \log P_t$ and $\Delta p_t = p_t - p_{t-1}$ and where the world interest rate in terms of traded goods is approximated as $\log(1 + r_t) \approx r_t$. Further, r_t^* can be interpreted as a consumption based real interest rate composed of the real interest rate and changes in the relative price of nontraded goods, see Bergin and Sheffrin (2000) for a detailed discussion. In the empirical estimation we demean r_t^* such that the constant drops out.

Defining net output as $NO_t \equiv Y_t - I_t - G_t$, the current account can be written as $CA_t = NO_t - C_t - r_t B_{t-1}$. Bergin and Sheffrin (2000) log-linearize the lifetime budget constraint of the household assuming a steady state where net foreign assets are zero. Combining it with the Euler equation (3.3.3) yields

$$no_t - c_t = -E_t \sum_{i=1}^{\infty} \beta^i \left[\Delta no_{t+i} - \gamma r_{t+i}^{*} \right], \tag{3.3.5}$$

with $no_t = \log NO_t$ and $\Delta no_t = no_t - no_{t-1}$. The left hand side of the above equation resembles the definition of the current account except that its components are in logs. Defining $CA_t^* \equiv no_t - c_t$ as the transformed representation of the current account, equation

[5] See Bergin and Sheffrin (2000) for a detailed derivation.

(3.3.5) can be rewritten as:

$$CA_t^* = -E_t \sum_{i=1}^{\infty} \beta^i \left[\Delta no_{t+i} - \gamma r_{t+i}^* \right]. \tag{3.3.6}$$

According to equation (3.3.6), the current account variable, CA_t^*, should include all of consumers' information on future changes of net output and of the consumption based interest rate. We test the validity of this equation in two different ways as it has been often done in the intertemporal approach. For the first test, we need to elicit consumers' expectations of future realizations of net output and the consumption based interst rate. For this purpose, we estimate the following VAR:

$$\begin{bmatrix} \Delta no \\ CA^* \\ r^* \end{bmatrix}_t = \begin{bmatrix} a_{11} & a_{12} & a_{13} \\ a_{21} & a_{22} & a_{23} \\ a_{31} & a_{32} & a_{33} \end{bmatrix} \begin{bmatrix} \Delta no \\ CA^* \\ r^* \end{bmatrix}_{t-1} + \begin{bmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \end{bmatrix}. \tag{3.3.7}$$

As argued by Ghosh (1995) the current account variable is included in the regression because under the null that the intertemporal model holds true, it contains all information of the household which the econometrician has no access to (e.g. expectations of shocks to government spending). The above VAR can be written more compactly as $\mathbf{z}_t = \mathbf{A}\mathbf{z}_{t-1} + \mathbf{u}_t$ with $z_t = [\Delta no_t, CA_t^*, r_t^*]'$ and where $\mathbf{u}_t$ is a mean zero vector of homoskedastic errors such that $E_t \mathbf{z}_{t+i} = \mathbf{A}^i \mathbf{z}_t$. Additional lags can be easily included in this notation by writing the VAR in the companion form. The VAR includes no constant since we demean all time series prior to estimation. Given consumers' forecasts of net output and the consumption based interest rate from the VAR as specified by equation (3.3.7), we can rewrite equation (3.3.6) as follows

$$\mathbf{h}\mathbf{z}_t = -E_t \sum_{i=1}^{\infty} \beta^i (\mathbf{g}_1 - \gamma \mathbf{g}_2) \mathbf{A}^i \mathbf{z}_t, \tag{3.3.8}$$

where $\mathbf{h} = [0\ 1\ 0]$, $\mathbf{g_1} = [1\ 0\ 0]$ and $\mathbf{g}_2 = [0\ 0\ 1]$. Defining the right hand side of equation (3.3.8) as the predicted current account variable, $\widehat{CA}_t^*$, we have that:

$$\widehat{CA}_t^* = \mathbf{k}\mathbf{z}_t, \tag{3.3.9}$$

where $\mathbf{k} = -(\mathbf{g_1} - \gamma \mathbf{g_2}) \beta \mathbf{A} (\mathbf{I} - \beta \mathbf{A})^{-1}$. If the intertemporal model holds true, then it must hold that $CA^* = \widehat{CA}_t^*$. Therefore, according to the intertemporal model the k-vector should be given by $\mathbf{k} = [0\,1\,0]$. This is a testable cross-equation restriction on the estimated VAR and dates back to Campbell (1987). Making use of the delta method, Bergin and Sheffrin (2000) show that the difference between the actual and the hypothesized value of $\mathbf{k}$ is χ^2-distributed with 3 degrees of freedom.[6] We call this the k-test. Below, we also plot $\widehat{CA}_t^*$ against CA_t^* for a visual comparison between the actual and the predicted current account variable allowing us to judge whether the German current account surplus is in line with predictions of the intertemporal model.

The second test also dates back to Campbell (1987). To derive this test we rewrite equation (3.3.6) as:

$$E_t CA_{t+1}^* = E_t \Delta no_{t+1} - \gamma E_t r_{t+1}^* + \frac{1}{\beta} CA_t^*. \tag{3.3.10}$$

Defining $R_t \equiv CA_t^* - (\Delta no_t - \gamma r_t^*) - (1/\beta) CA_{t-1}^*$, equation (3.3.10) implies that R_t should be unpredictable given past information. Formally, it should hold that $E_t(R_t | I_{t-1}) = 0$, where I_{t-1} summarizes all information up to and including period $t-1$. Hence, variable R_t should be unpredictable according to the intertemporal model. We call this the R-test, which we conduct by regressing R_t on past values of CA_t^*, Δno_t and r_t^* and by testing the joint nullity of the coefficients.

3.4 Data and parameter values

Our empirical analysis is based on quarterly data which are seasonally adjusted at annual rates. Because of the German reunification our sample starts in 1991-Q1 and runs until 2017-Q4. The data are retrieved from the International Financial Statistics (IFS), published by the International Monetary Fund (IMF), unless otherwise stated. We construct the series for net output by subtracting government expenditure (G_t) and investment expenditure (I_t) from GDP (Y_t). The current account variable (CA_t^*) is defined as the difference of log net output and the log of private consumption expenditure (C_t). All variables are adjusted by the GDP-deflator and are expressed as per-capita ratios.

The consumption based interest rate (r^*) is calculated according to (3.3.4). We compute the real world interest rate (r_t) following Barro and Sala-i-Martin (1990). That is, we adjust

[6] Below we also test a simpler version of the intertemporal model without time varying interest rates. In this case and when we consider further lags in the VAR the k-vector and its components are adjusted appropriately as is the number of degrees of freedom.

the short term nominal interest rates of the G-7 economies by inflation expectations in order to get an "ex-ante" real interest rate. Short term nominal interest rates are from the OECD database because of better data availability. For Japan, however, we take the IFS Treasury Bill rate also due to data availability. Inflation expectations in each country are estimated by a six-quarter autoregression using the respective country's consumer price index.[7] In order to compute the world real interest rate we weight each country-specific real rate by its time-varying share of real GDP in the G-7.[8]

The second component required to compute the consumption based interest rate, r^*, is a measure of the ex-ante expectation of P_t. Following Bergin and Sheffrin (2000) and Rogoff (1992), we use the real exchange rate as a proxy for P_t. For this purpose, we take the real effective exchange rate based on Germany's Consumer Price Index from the IFS. We estimate a six-quarter autoregression of this series and take logs and first differences in order to compute the ex-ante expected exchange rate appreciation/depreciation ($E_{t-1}\Delta p_t$). Finally, as mentioned above the series for Δno, CA^* and r^* are all demeaned.

We further need to assign parameter values. The discount factor β is computed as $\beta = 1/(1+\bar{r})$, where $\bar{r}$ is the average world interest rate.[9] For our sample we get $\beta = 0.99$. For the share of traded goods in private final consumption we take the estimate from Campa and Gavilan (2011) for Germany of $a = 0.36$ which is estimated from the input-output information provided by Eurostat with data from 1995. In a robustness analysis we also consider a lower value for a because the share of traded-goods in consumption is commonly expected to have declined in developed countries due to the growth of the service sector.

Finally, we have to assign a value for the intertemporal elasticity of substitution γ. As we will show below, our results are sensitive to the choice of γ. Following Bergin and Sheffrin (2000), we choose γ in our baseline specification such that the variance of the predicted current account variable, $\widehat{CA}^*$, matches the variance of its counterpart in the data, CA^*.[10] For our baseline specification we get $\gamma = 0.047$ which is in line with Hall (1988). In a robustness check we choose $\gamma = 0.4$ in order to accommodate estimates for the coefficient of relative risk aversion σ (Mehra and Prescott, 1985). In another specification,

[7]The CPI-series start in 1989 in order to determine inflation expectations for our first sample period in 1991-Q1. For Germany the CPI series starts only in 1991-Q1, which is why we omit Germany in the computation of the world interest rate until 1992-Q3.

[8]The real GDP data are retrieved from the Worldbank's World Development Indicators.

[9]Sachs (1982) allows the discount factor to differ from $1/(1+\bar{r})$. In that case there is a "consumption-tilting" motive: depending on time preferences relative to the world interest rate consumers move consumption to the present or to the future, see also Ghosh (1995).

[10]This, however, reduces the degrees of freedom of the χ^2-statistic of the k-test and of the R-test by one.

Bergin and Sheffrin (2000) choose γ to minimize the χ^2-statistic of the k-test. In our case this yields a negative value for γ such that we refrain from this specification.

Before estimating the VAR we test whether the time series are stationary. We find that both the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) test reject the presence of a unit root for the net output series (logged and in first differences) and the consumption based interest rate, see Table 3.1. In line with these results, the KPSS test does not reject the null of stationarity for both series. The ADF and the PP test, however, do not reject the hypothesis that a unit root is present in the series for the current account variable. The KPSS test further rejects the null of stationarity. The failure to reject a unit root for the current account is a frequent result in the literature testing the intertemporal model.[11] It is commonly argued that the failure to reject the unit root results from low power of these tests in the borderline case of a highly persistent stationary processes (see also Cochrane, 1991; Blough, 1992). Similarly, in small samples the KPSS test often spuriously rejects the null of stationarity when faced with data from a highly persistent stationary processes (Caner and Kilian, 2001).

Economic theory also provides strong reasons for why the current account should be stationary. In our model, for instance, the current account variable is stationary if net output (in first differences) and the consumption based interest rate are stationary, see equation (3.3.6). Further, since the current account variable is defined as $CA_t^* = \log(NO_t) - \log(C_t)$, a non-stationary current account variable would imply a divergence of net output from consumption. On a balanced growth path the ratio of net output to consumption should, however, be constant such that the series for the current account variable should be stationary. More specifically, our model is consistent with output growing with a linear trend subject to permanent shocks (e.g. due to technological progress). In this case the change in net output consists of the trend and the permanent shock. Since we demean all variables prior to estimating the VAR, the currrent account variable as determined by (3.3.6) is therefore adjusted for trend growth. For these reasons, we consider the current account as stationary but highly persistent in our sample.[12]

However, the high persistence of the current account might be problematic for the k-test. The reason is that the delta method used to compute the χ^2-statistic can be inaccurate and

[11] See for instance Sheffrin and Woo (1990), Huang (1993), Gruber (2004), Campa and Gavilan (2011) and Bussiere et al. (2018).

[12] Other theoretical models in which the current account of a small open economy is stationary include: Galí and Monacelli (2005a) and Adolfson et al. (2008).

Table 3.1: Unit root tests

Number of lags	1	3	5
Change in net output (Δno)			
ADF	-7.795***	-5.441***	-3.91***
PP	-12.545***	-12.625***	-12.661***
KPSS	0.077	0.090	0.095
Current account (CA^*)			
ADF	-0.875	-1.004	-0.731
PP	-1.305	-1.263	-1.223
KPSS	4.59***	2.38***	1.63***
Consumption based real interest rate (r^*)			
ADF	-5.790***	-5.377***	-3.933***
PP	-9.131***	-9.159***	-9.122***
KPSS	0.107	0.095	0.103

Notes: range is 1992-Q4 to 2017-Q4. ADF is the augmented Dickey-Fuller test (number of lags refers to differenced term); PP is the Phillips-Perron test; KPSS is the Kwiatkowski-Phillips-Schmidt-Shin test (H_0: stationarity). Time series are not demeaned, all tests include a constant but no time trend. *, ** and *** indicate significance at the 10%, 5% and 1% level. In calculating r^* we chose $a = 0.36$ and $\gamma = 0.047$.

could lead to over-rejection but also to over-acceptance of the intertemporal model. For this reason we prefer the R-test which avoids such problems (see Miniane and Mercereau, 2004, for both points). This is also confirmed in Bouakez and Kano (2009) who conduct a Monte Carlo study within the simple intertemporal model using UK data. They show that the R-test has the appropriate size while the k-test is biased towards over-rejecting. In Section 3.6 we also conduct a Monte Carlo analysis to analyze whether our findings are consistent with the intertemporal model when it cannot be rejected that a unit root is present in the time series for the current account. We confirm that the R-test has the appropriate size. Contrary to the results in Bouakez and Kano (2009), however, in our Monte Carlo analysis the k-test has a small bias to over-accept the cross-equation restrictions of the intertemporal model.

3.5 Results

Table 3.2 displays our results for the tests of the intertemporal model, namely the k-test and the R-test as described in Section 3.3. The results are based on a VAR with one lag.

Table 3.2: Results of k-test and R-test

	Baseline	Alternative specifications			
	(1) γ chosen to match variance with a = 0.36	(2) Simple model r* constant	(3) higher elast. of intertemp. subst. $\gamma = 0.4$	(4) γ chosen to match variance with a = 0.25	(5) just interest rate, exchange rate excluded
γ	0.047	-	0.4	0.056	0.047
k-vector					
no_t	0.115	0.092	0.396	0.121	0.122
	(0.146)	(0.060)	(0.296)	(0.163)	(0.062)
CA_t^*	1.001	0.909	-1.887	0.998	0.572
	(0.930)	(0.375)	(1.852)	(1.039)	(0.387)
r_t^*	0.010	-	0.190	0.011	0.084
	(0.007)		(0.172)	(0.009)	(0.096)
χ^2-statistic	3.64	4.17	3.98	3.26	4.73
p-val. k-test	$0.162^\dagger$	0.125	0.264	$0.196^\dagger$	0.193
p-val. R-test	0.905	0.185	0.000	0.953	0.234
$\sigma_{\widehat{CA}^*}/\sigma_{CA^*}$	1.00	0.91	1.94	1.00	0.55

Notes: Standard errors in parentheses. Regressions are for 1991-Q1 to 2017-Q4. Share of tradeables in consumption, a, is 0.36, unless otherwise stated. $\beta = 0.99$. † indicates degrees of freedom equal to 2 instead of 3, as γ chosen to match the variance of CA^*.

This lag length is suggested by the AIC criterion.[13] Each column represents a different model specification and reports the estimated k-vector, the associated χ^2-statistic of the k-test and its p-value. Further, it reports the p-value of the R-test and the ratio of the standard deviation of the predicted current account variable, $\widehat{CA}^*$, to its counterpart in the data, CA^*.

The first column is the most important one. It reports the results for the baseline specification in which the intertemporal elasticity of substitution is chosen such that the variance of $\widehat{CA}^*$ equals the variance of CA^* (as discussed in Section 3.4). Under this specification neither the k-test nor the R-test reject the cross-equation restrictions of the intertemporal model. Hence, for the baseline we find that the German current account data are consistent with the intertemporal model. Further, Figure 3.2 displays the current account variable as predicted by the cross-equation restriction (3.3.9), that is $\widehat{CA}^*$, and

[13] Our results are also robust to including more lags. In this case, however, parameter γ has to be adjusted as discussed in Section 3.4.

its counterpart in the data, CA^*. The predicted current account variable matches the data strikingly well—also for the recent period with the large surplus. Hence, the German current account surplus appears not too large relative to the prediction of the intertemporal model.

In order to understand whether expectations about net output or the consumption based interest rate drive the predicted current account variable, we decompose $\widehat{CA}^*$ into both components. Formally, equations (3.3.6) and (3.3.9) imply:

$$1 = -E_t \sum_{i=1}^{\infty} \beta^i \frac{\Delta\widehat{no}_{t+i}}{CA_t^*} + \gamma E_t \sum_{i=1}^{\infty} \beta^i \frac{\widehat{r}_{t+i}^*}{CA_t^*}$$

We find that on average 96% of the predicted current account are due to expectations of changes in net output. This number changes only marginally, if we only consider the average contribution of net output between 2011 and 2017. In this case net output contributes 95% of the predicted current account variable. In other words, changes in expectations over net output play the dominant role in explaining the German current account surplus.

Columns 2-5 of Table 3.2 report results for the k-test and the R-test for alternative model specifications. In column 2 we report results for the simple intertemporal model where the consumption based real interest rate, r^*, is constant. Again, neither the k-test nor R-test reject the intertemporal model. The upper-left panel of Figure 3.3 shows that the predicted current account variable fits the actual current account variable again very well. This is in line with the previous result that a time-varying interest rate contributes relatively little in explaining the German current account.

In column 3 of Table 3.2 we choose a higher value for the intertemporal elasticity of substitution with $\gamma = 0.4$. The R-test clearly rejects the model while the k-test is far from rejecting the model. The upper-right panel in Figure 3.3 also indicates the poor fit of the model under this specification.[14] Overall we find that the fit of the model is very sensitive to the choice of γ. However, since the model with constant interest rates fits the current account very well, we do not consider this sensitivity as problematic.

As a further robustness check, we choose a lower value for the share of tradeables in consumption with $a = 0.25$ and report results in column 4 of Table 3.2. In this exercise we also change the value for γ such that the variance of the predicted current account variable

[14]This finding strengthens our view that the R-test is better suited to infer the validity of the model, see the discussion at the end of section 3.4.

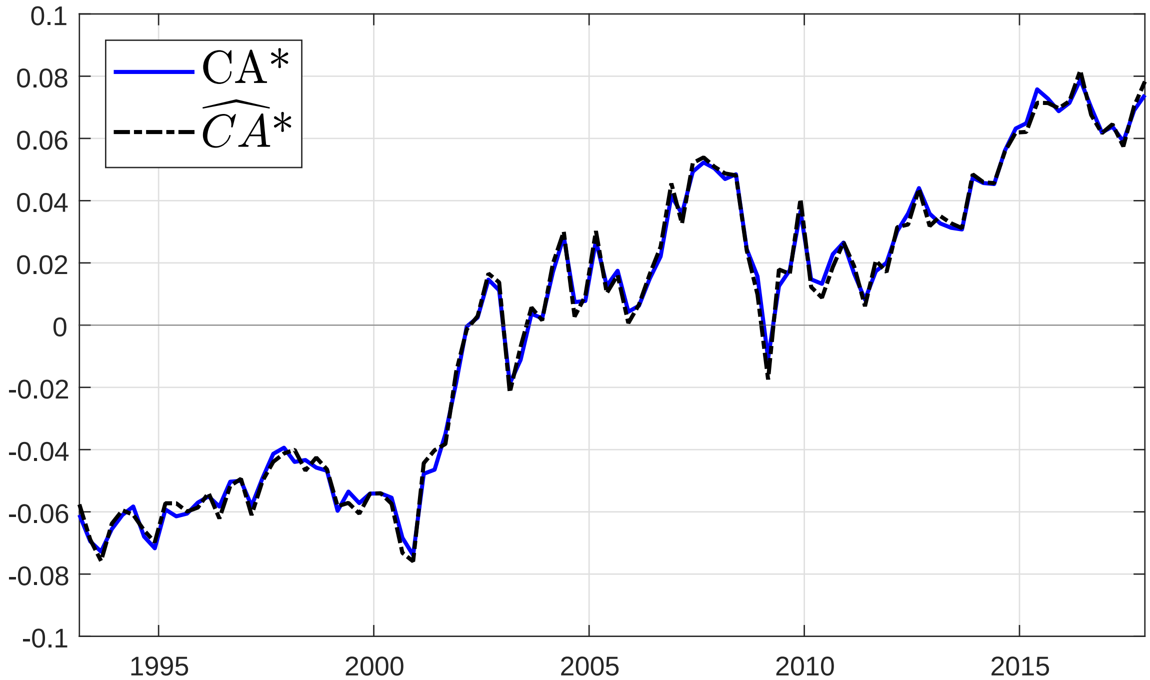

Figure 3.2: Actual German current account variable, CA^*, and as predicted by the cross-equation restrictions of the intertemporal model, $\widehat{CA}^*$, under the baseline parameterization (see column 1 in Table 3.2).

$\widehat{CA}^*$ matches the variance of CA^* (such that this exercise is comparable to column 1 in Table 3.2). Overall these variations have little effect on the test statistics. Also the fit of the predicted current account appears to be unaffected by these changes as can be seen in the lower-left panel of Figure 3.3. Results are also robust to higher values of a (not shown).

Finally, as in Bergin and Sheffrin (2000) we consider the case where changes in the consumption based real interest rate are only due to changes in the world interest rate and not due to changes in the relative price of tradables (proxied by the exchange rate as discussed in section 3.4). Results are reported in column 5 of Table 3.2. Importantly, we keep parameters otherwise as in the baseline. Again, the k-test and the R-test do not reject the cross equation restrictions on the k-vector. The lower-right panel of Figure 3.3 shows, however, that the fit of the predicted current account variable worsens. Therefore, the exchange rate appears to be more important in explaining the current account variable compared to the real world interest rate.

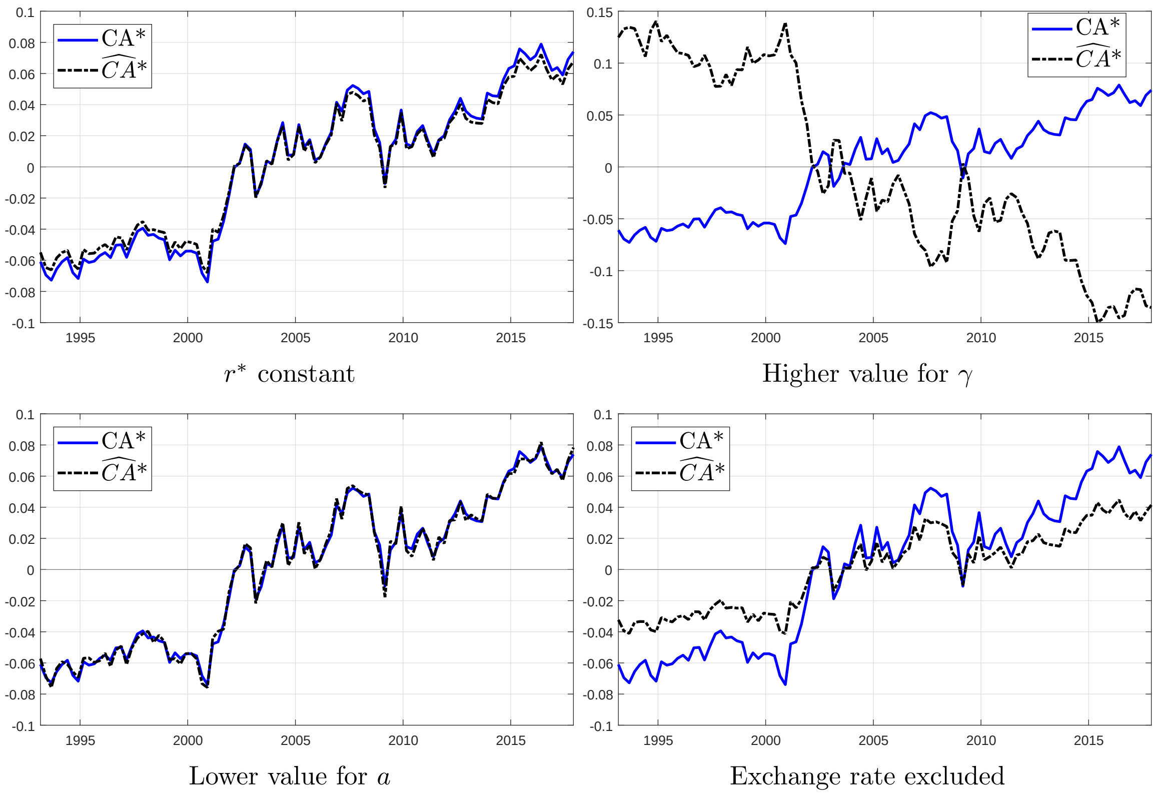

Figure 3.3: Actual current account variable, CA^*, and as predicted by the model, $\widehat{CA}^*$, for different model specifications (see columns 2-5 in Table 3.2).

3.6 Monte Carlo analysis

Given the good fit of the current account variable as displayed in Figure 3.2, a concern is that our results are driven by a potential unit root which we cannot reject in the time series for the current account variable. To assess this possibility we conduct a Monte Carlo analysis. For this purpose we simulate data from a *stationary* intertemporal model and consider a specification for which we cannot reject that a unit root is present in the simulated series for the current account. We then perform the k-test, the R-test and compare the graphical fit of the predicted current account to the simulated time series.

For our Monte Carlo analysis we resort to the small open economy model with a debt elastic interest rate put forward in Schmitt-Grohé and Uribe (2003). The reason to deviate from the model of Bergin and Sheffrin (2000) is that their model is in partial equilibrium.

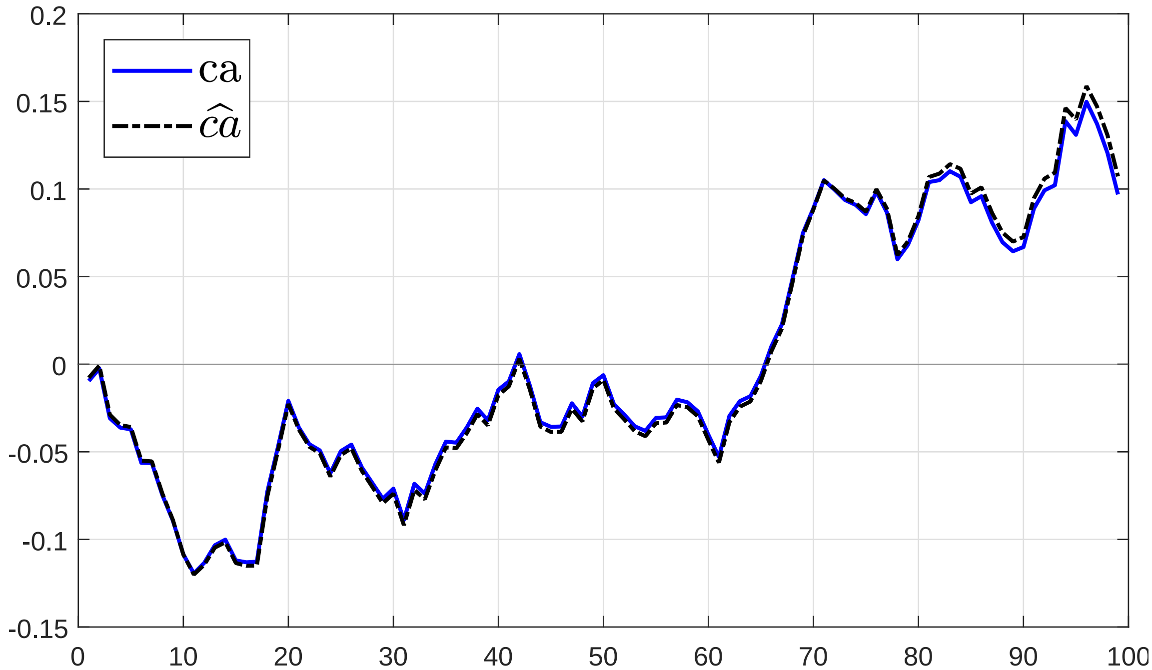

Figure 3.4: Example of the good graphical fit of the current account within our Monte Carlo analysis for a simulation of 100 periods (as in our empirical analysis). Simulated current account series, ca (solid line), and as predicted by the cross-equation restrictions, $\widehat{ca}$ (dashed line).

To avoid stochastic singularity and to increase the autocorrelation of the current account we introduce a shock to the discount factor and a labor supply shock to the model of Schmitt-Grohé and Uribe. The Appendix describes the model and its calibration in detail.

We solve the model and simulate the time series using Dynare (Adjemian et al., 2011). Each simulated time series contains 100 observations (as in our empirical setting) after dropping the initial 200 periods as a burn-in phase. We conduct the ADF test on the simulated current account series and consider 0 to 5 lags of the differenced term when computing the ADF test. We keep the simulated time series, if the ADF test does not reject the presence of a unit root at all specified lags (as in our empirical setting). Otherwise we discard the simulation. We continue with this procedure until we retained 2000 simulations. On these simulated time series we perform the k-test, the R-test and compare the graphical fit of the current account to the one predicted by the cross-equation restriction.

We find that the R-test has the appropriate size on the simulated data (5.6% at the 5%

significance level) while the k-test tends to over-accept the cross-equation restrictions of the intertemporal model (the rejection rate is 3.2% at the 5% significance level). We further find that the predicted current account fits the simulated current account very well—as is the case in our empirical analysis. Figure 3.4 shows an example of the good graphical fit (which is a general result). Overall, the Monte Carlo shows that our empirical findings are in line with a stationary intertemporal model and that our tests and graphical comparisons are working well when the presence of a unit root cannot be rejected.

3.7 Conclusion

It is often argued that the German current account surplus is too large (as we discuss in our introduction). In this paper we investigate this issue through the lens of the intertemporal approach to the current account. We find that the intertermporal model is not rejected for Germany and that the current account surplus is in line with model predictions. We further showed that the surplus can be rationalized as an increase in savings in response to expectations of relatively lower future income.

When estimating the intertemporal model we assumed that the time series for the current account is stationary even though standard statistical tests rejected stationarity or could not reject the presence of a unit root. We justified our assumption by the difficulty of these tests to distinguish between the borderline case of a highly persistent process and a unit root and by economic theory which suggests that the current account should be stationary. We further conducted a Monte Carlo analysis to assess the possibility that our results are driven by a potential unit root. In this exercise we generated highly persistent current account data from a stationary intertemporal model with the feature that it could not be rejected that the simulated data are non-stationary. Our Monte Carlo analysis showed that our empirical results are in line with the intertemporal model and hence that the failure to reject a unit root is not per se problematic.

However, we stress that even if a current account surplus (or deficit) can be justified by households' expectations within the intertemporal model, there is no guarantee that these expectations materialize. As Campa and Gavilan (2011) show in their analysis of the intertemporal model, Spain's expectation of future income growth were at its height just before the financial crisis hit. Further, in order to perform the present value tests of the intertemporal model one has to compute households' expectations of changes in net output over the infinite horizon. These expectations are modeled by a VAR. It is well known, that errors due to potential misspecifications of the VAR cumulate at the forecast horizon (see, for instance, Jordà, 2005).

Given these caveats, we showed that the large German current account surplus can be rationalized within the intertemporal approach to the current account. Hence, our results put the view into question that the German current account surplus is too large.

3.A The model underlying the Monte Carlo analysis

The underlying model of our Monte Carlo analysis is based on the small open economy model with a debt elastic interest rate in Schmitt-Grohé and Uribe (2003). We modify this model by considering separable utility between consumption and labor and by introducing a shock to the discount factor (b_t) and a labor supply shock (χ_t). The representative household of the small open economy has the objective to maximize expected lifetime utility which is given by:

$$E_0 \sum_{t=0}^{\infty} \beta^t b_t \left(\frac{c_t^{1-\sigma}}{1-\sigma} + \omega \chi_t \log(1-h_t) \right) \tag{A.1}$$

where consumption is denoted by c_t and labor by h_t. Further $\sigma, \omega > 0$ and β is the subjective discount factor. Optimization is subject to the following law of motion for foreign debt, d_t:

$$d_t = (1 + r_{t-1})d_{t-1} - y_t + c_t + i_t + \frac{\Phi}{2}(k_{t+1} - k_t)^2 \tag{A.2}$$

where r_t is the domestic interest rate which is composed of an exogenous constant world interest rate r and a country-specific interest rate premium $p(d_t) = \psi(\exp(d_t - \bar{d}) - 1)$ with $\bar{d}$ denoting steady state debt. Put differently,

$$r_t = r + p(d_t) \tag{A.3}$$

We assume that the household treats the interest rate premium as exogenous. The purpose of this premium is to render the model stationary. Further, in equation (A.2) y_t denotes income with $y_t = A_t k_t^{\alpha} h_t^{1-\alpha}$ where k_t is capital and A_t is a technology shock whose stochastic process is specified below. Investment is denoted by i_t with $i_t = k_{t+1} - (1-\delta)k_t$ where $0 < \delta < 1$ is the depreciation rate. The last term in equation (A.2) represents quadratic capital adjustment costs with $\Phi > 0$. Substituting for y_t and i_t in equation (A.2) we get:

$$d_t = (1 + r_{t-1})d_{t-1} - A_t k_t^{\alpha} h_t^{1-\alpha} + c_t + k_{t+1} - (1-\delta)k_t + \frac{\Phi}{2}(k_{t+1} - k_t)^2 \tag{A.4}$$

Maximizing lifetime utility, (A.1), over d_t, c_t, h_t and k_{t+1} subject to equation (A.4) yields the following first order conditions:

$$b_t \lambda_t = \beta E_t b_{t+1} \lambda_{t+1}(1 + r_t) \tag{A.5}$$

$$c_t^{-\sigma} = -\lambda_t \tag{A.6}$$

$$\omega \chi_t \frac{1}{1-h_t} = c_t^{-\sigma} A_t (1-\alpha) \left(\frac{k_t}{h_t}\right)^{\alpha} \tag{A.7}$$

$$b_t \lambda_t \left(1 + \Phi(k_{t+1} - k_t)\right) = \beta E_t b_{t+1} \lambda_{t+1} \left(A_{t+1} \alpha \left(\frac{k_{t+1}}{h_{t+1}}\right)^{\alpha - 1} + (1-\delta) + \Phi(k_{t+2} - k_{t+1}) \right) \tag{A.8}$$

where λ_t is the Lagrange multiplier. For the shock processes we assume that:

$$\log A_t = \rho_a \log A_{t-1} + \sigma_a \varepsilon_{a,t} \tag{A.9}$$

$$\log b_t = \rho_b \log b_{t-1} + \sigma_b \varepsilon_{b,t} \tag{A.10}$$

$$\log \chi_t = \rho_\chi \log \chi_{t-1} + \sigma_\chi \varepsilon_{\chi,t} \tag{A.11}$$

where $\rho_a, \rho_b, \rho_\chi \in (0,1)$ and $\varepsilon_{a,t}$, $\varepsilon_{b,t}, \varepsilon_{\chi,t} \sim iid\mathcal{N}(0,1)$. The innovations $\varepsilon_{a,t}$, $\varepsilon_{b,t}$, and $\varepsilon_{\chi,t}$ are assumed to be uncorrelated at all leads and lags.

A competitive equilibrium is defined as a collection of stochastic processes $\{c_t, h_t, d_t, r_t, k_t, \lambda_t\}_{t=0}^{\infty}$ given initial values d_{-1}, k_0 and equations (A.9) – (A.11) such that equations (A.3) – (A.8) are fulfilled. Once we solved the model, the current account can be calculated by computing the change in net foreign assets:

$$ca_t = d_{t-1} - d_t$$

We largely follow Schmitt-Grohé and Uribe (2003) in calibrating the model. The calibration is summarized in Table 3.3. The time unit of the model corresponds to one year. We solve the model and simulate the time series using Dynare (Adjemian et al., 2011). Each simulated time series contains 100 observations (as in our empirical setting) after dropping

Table 3.3: Parameter values

r	0.04	As in Schmitt-Grohé and Uribe (2003)
β	0.9615	Discount factor, implied by $\beta(1+r)=1$
σ	9	Low elasticity of intertemporal substitution as in Section 3.4
ω	16.42	Calibrated such that the household works 1/3 of her time in steady state
Φ	0.28	Set to get a high autocorrelation of the current account
ψ	0.0000742	Slightly lower compared to Schmitt-Grohé and Uribe (2003) to increase the autocorrelation of the current account
$\bar{d}$	0	As in Bergin and Sheffrin (2000)
α	0.32	As in Schmitt-Grohé and Uribe (2003)
δ	0.01	As in Schmitt-Grohé and Uribe (2003)
ρ_a	0.75	Set to get a high autocorrelation of the current account
σ_a	0.0129	As in Schmitt-Grohé and Uribe (2003)
ρ_b	0.95	Set to get a high autocorrelation of the current account
σ_b	0.029	Set to get a high autocorrelation of the current account
ρ_χ	0.95	Set to get a high autocorrelation of the current account
σ_χ	0.059	Set to get a high autocorrelation of the current account

the initial 200 periods as a burn-in phase. We conduct the ADF test on the simulated current account series and consider 0 to 5 lags of the differenced term when computing the ADF test. We keep the simulated time series, if the ADF test does not reject the presence of a unit root at all specified lags (as in our empirical setting). Otherwise we discard the simulation. We continue with this procedure until we retained 2000 simulations. Given the simulated time series, we perform the k-test and the R-test as described in Section 3.3. We calculate these tests with the values for β and $\gamma = 1/\sigma$ which were used in the simulation (see Table 3.3). As described in Section 3.6, we find that the R-test has the appropriate size on the simulated data (5.6% at the 5% significance level) while the k-test tends to over-accept the cross equation restrictions of the intertemporal model (the rejection rate is 3.2% at the 5% significance level).[15]

[15] In our Monte Carlo exercise we apply the k- and the R-test as introduced in Section 3.3. As previously mentioned, the model in our Monte Carlo analysis differs from the one in Bergin and Sheffrin (2000) and therefore implies a slightly different cross-equation restriction. Since the size of the R-test is correct and the one of the k-test has only a minor bias, using the cross-equation restriction as described in Section 3.3 seems therefore unproblematic given our calibration.

References

Acconcia, Antonio, Giancarlo Corsetti, and Saverio Simonelli, “Mafia and Public Spending: Evidence on the Fiscal Multiplier from a Quasi-experiment,” *American Economic Review*, 2014, *104* (7), 2185–2209.

Adam, Klaus and Roberto M. Billi, “Optimal Monetary Policy under Commitment with a Zero Bound on Nominal Interest Rates,” *Journal of Money, Credit and Banking*, October 2006, *38* (7), 1877–1905.

_ and _ , “Discretionary monetary policy and the zero lower bound on nominal interest rates,” *Journal of Monetary Economics*, April 2007, *54* (3), 728–752.

_ , Pei Kuang, and Albert Marcet, “House Price Booms and the Current Account,” *NBER Macroeconomics Annual*, 2012, *26* (1), 77–122.

Adjemian, Stéphane, Houtan Bastani, Michel Juillard, Ferhat Mihoubi, George Perendia, Marco Ratto, and Sébastien Villemot, “Dynare: Reference manual, version 4,” 2011.

Adolfson, Malin, Stefan Laséen, Jesper Lindé, and Mattias Villani, “Evaluating an estimated new Keynesian small open economy model,” *Journal of Economic Dynamics and Control*, 2008, *32* (8), 2690–2721.

Ahmed, Shaghil, “Temporary and permanent government spending in an open economy: some evidence for the United Kingdom,” *Journal of Monetary economics*, 1986, *17* (2), 197–224.

_ and John H Rogers, “Government budget deficits and trade deficits Are present value constraints satisfied in long-term data?,” *Journal of Monetary Economics*, 1995, *36* (2), 351–374.

Bacchetta, Philippe and Eric van Wincoop, “Infrequent Portfolio Decisions: A Solution to the Forward Discount Puzzle,” *American Economic Review*, June 2010, *100* (3), 870–904.

Backus, David K. and Gregor W. Smith, "Consumption and real exchange rates in dynamic economies with non-traded goods," *Journal of International Economics*, November 1993, *35* (3-4), 297–316.

Barro, Robert J. and Xavier Sala-i-Martin, "World Real Interest Rates," *NBER Macroeconomics Annual*, 1990, *5*, 15–61.

Beetsma, Roel and Harald Uhlig, "An analysis of the stability and growth pact," *Economic Journal*, 1999, (109), 546–571.

_ **and Henrik Jensen**, "Monetary and fiscal policy interactions in a micro-founded model of a monetary union," *Journal of International Economics*, 2005, (67), 320–352.

Beetsma, Roel M. W. J. and A. Lans Bovenberg, "Monetary union without fiscal coordination may discipline policymakers," *Journal of International Economics*, August 1998, *45* (2), 239–258.

Benigno, Gianluca and Pierpaolo Benigno, "Price Stability in Open Economies," *Review of Economic Studies*, October 2003, *70* (4), 743–764.

_ **and** _ , "Designing targeting rules for international monetary policy cooperation," *Journal of Monetary Economics*, April 2006, *53* (3), 473–506.

Benigno, Pierpaolo, "Optimal monetary policy in a currency area," *Journal of International Economics*, 2004, *63* (2), 293 – 320.

_ **and Michael Woodford**, "Optimal taxation in an RBC model: A linear-quadratic approach," *Journal of Economic Dynamics and Control*, 2006, *30* (9-10), 1445–1489.

Bergin, Paul R and Steven M Sheffrin, "Interest Rates, Exchange Rates and Present Value Models of the Current Account," *Economic Journal*, April 2000, *110* (463), 535–558.

Bjornland, Hilde C., "Monetary policy and exchange rate overshooting: Dornbusch was right after all," *Journal of International Economics*, 2009, *79* (1), 64 – 77.

Blanchard, Olivier, Christopher Erceg, and Jesper Lindé, "Jump-Starting the Euro Area Recovery: Would a Rise in Core Fiscal Spending Help the Periphery?," in "NBER Macroeconomics Annual 2016, Volume 31" NBER Chapters, National Bureau of Economic Research, Inc, March 2016.

Blough, Stephen R., "The relationship between power and level for generic unit root tests in finite samples," *Journal of Applied Econometrics*, 1992, *7* (3), 295–308.

Born, Benjamin, Falko Juessen, and Gernot J. Müller, "Exchange rate regimes and fiscal multipliers," *Journal of Economic Dynamics and Control*, 2013, *37* (2), 446–465.

Bouakez, Hafedh and Michel Normandin, "Fluctuations in the foreign exchange market: How important are monetary policy shocks?," *Journal of International Economics*, 2010, *81* (1), 139–153.

_ **and Takashi Kano**, "Tests of the present-value model of the current account: a note," *Applied Economics Letters*, 2009, *16* (12), 1215–1219.

Bruno, Valentina and Hyun Song Shin, "Capital flows and the risk-taking channel of monetary policy," *Journal of Monetary Economics*, 2015, *71*, 119–132.

Bussiere, Matthieu, Karadimitropoulou Aikaterini, and Miguel A. Leon-Ledesma, "Current Account Dynamics and the Real Exchange Rate: Disentangling the Evidence," *Macroeconomic Dynamics*, July 2018, p. 26.

Bussière, Matthieu, Marcel Fratzscher, and Gernot J. Müller, "Productivity shocks, budget deficits and the current account," *Journal of International Money and Finance*, 2010, *29* (8), 1562 – 1579.

Calvo, Guillermo A., "Staggered Prices in a Utility-Maximizing Framework.," *Journal of Monetary Economics*, 1983, *12* (3), 383 – 398.

Campa, Jose Manuel and Angel Gavilan, "Current accounts in the euro area: An intertemporal approach," *Journal of International Money and Finance*, 2011, *30* (1), 205 – 228.

Campbell, John Y, "Does Saving Anticipate Declining Labor Income? An Alternative Test of the Permanent Income Hypothesis," *Econometrica*, November 1987, *55* (6), 1249–1273.

_ **and Robert J Shiller**, "Cointegration and tests of present value models," *Journal of political economy*, 1987, *95* (5), 1062–1088.

Caner, Mehmet and Lutz Kilian, "Size distortions of tests of the null hypothesis of stationarity: evidence and implications for the PPP debate," *Journal of International Money and Finance*, 2001, *20* (5), 639–657.

Chari, V. V. and Patrick J. Kehoe, "International Coordination of Fiscal Policy in Limiting Economies," *Journal of Political Economy*, 1990, *98* (3), 617–636.

Chetty, Raj, Adam Guren, Day Manoli, and Andrea Weber, "Are Micro and Macro Labor Supply Elasticities Consistent? A Review of Evidence on the Intensive and Extensive Margins," *American Economic Review*, May 2011, *101* (3), 471–475.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans, "Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy," *Journal of Political Economy*, 2005, *113* (1), 1–45.

Christiano, Lawrence, Martin Eichenbaum, and Sergio Rebelo, "When Is the Government Spending Multiplier Large?," *Journal of Political Economy*, 2011, *119* (1), 78 – 121.

Clarida, Richard, Jordi Gali, and Mark Gertler, "A simple framework for international monetary policy analysis," *Journal of Monetary Economics*, July 2002, *49* (5), 879–904.

Cochrane, John H., "A critique of the application of unit root tests," *Journal of Economic Dynamics and Control*, 1991, *15* (2), 275 – 284.

Coibion, Olivier and Yuriy Gorodnichenko, "What Can Survey Forecasts Tell Us about Information Rigidities?," *Journal of Political Economy*, 2012, *120* (1), 116–159.

_ and _ , "Information Rigidity and the Expectations Formation Process: A Simple Framework and New Facts," *American Economic Review*, August 2015, *105* (8), 2644–78.

_ , _ , Lorenz Kueng, and John Silvia, "Innocent Bystanders? Monetary policy and inequality," *Journal of Monetary Economics*, 2017, *88* (C), 70–89.

Cook, David and Michael B. Devereux, "Optimal fiscal policy in a world liquidity trap," *European Economic Review*, May 2011, *55* (4), 443–462.

_ and _ , "Sharing the Burden: Monetary and Fiscal Responses to a World Liquidity Trap," *American Economic Journal: Macroeconomics*, July 2013, *5* (3), 190–228.

Corsetti, Giancarlo and Paolo Pesenti, "Welfare And Macroeconomic Interdependence," *The Quarterly Journal of Economics*, May 2001, *116* (2), 421–445.

_ , André Meier, Keith Kuester, and Gernot J. Müller, "Sovereign risk and belief-driven fluctuations in the euro area," *Journal of Monetary Economics*, 2014, (61), 53–73.

_ , Keith Kuester, and Gernot J. Müller, *Floats, pegs and the transmission of fiscal policy.*, Vol. 17 of *Fiscal Policy and Macroeconomic Performance, Central Banking, Analysis, and Economic Policies Book Series*, Santiago, Chile, Central Bank of Chile: in Céspedes, L.F. and J. Galí (eds.),

_ , _ , and _ , "Fixed on Flexible: Rethinking Exchange Rate Regimes after the Great Recession," *IMF Economic Review*, August 2017, *65* (3), 586–632.

_ , _ , André Meier, and Gernot J. Müller, "Sovereign Risk, Fiscal Policy, and Macroeconomic Stability," *The Economic Journal*, 2013, *123* (566), F99–F132.

_ , _ , André Meier, and Gernot J. Müller, "Debt Consolidation and Fiscal Stabilization of Deep Recessions," *American Economic Review*, May 2010, *100* (2), 41–45.

Cotarelli, Carlo, "Fiscal adjustment: Too much of a good thing?," 2012. Vox EU.

Cúrdia, Vasco and Michael Woodford, "The Central-Bank Balance Sheet as an Instrument of Monetary Policy," *Journal of Monetary Economics*, 2011, (58), 54–79.

Cúrdia, Vasco and Michael Woodford, "Credit Frictions and Optimal Monetary Policy," *Journal of Monetary Economics*, 2016, *84* (C), 30–65.

Cwik, Tobias and Volker Wieland, "Keynesian Government Spending Multipliers and Spillovers in the Euro Area," *Economic Policy*, 2011, (26), 493–549.

Devereux, Michael B., "The terms of trade and the international coordination of fiscal policy," *Econmic Inquiry*, 1991, (102(3)), 720–736.

Dixit, Avinash and Luisa Lambertini, "Monetary-fiscal policy interactions and commitment versus discretion in a monetary union," *European Economic Review*, May 2001, *45* (4-6), 977–987.

_ **and** _ , "Symbiosis of monetary and fiscal policies in a monetary union," *Journal of International Economics*, August 2003, *60* (2), 235–247.

Dornbusch, Rudiger, "Expectations and Exchange Rate Dynamics," *Journal of Political Economy*, December 1976, *84* (6), 1161–1176.

Eggertsson, Gauti B., "What Fiscal Policy is Effective at Zero Interest Rates?," in "NBER Macroeconomics Annual 2010, Volume 25" NBER Chapters, National Bureau of Economic Research, Inc, 2011, pp. 59–112.

_ **and Michael Woodford**, "The Zero Interest-Rate Bound and Optimal Monetary Policy," *Brookings Papers on Economic Activity*, 2003, *1*, 139–211.

_ **and Sanjay R. Singh**, "Log-linear Approximation versus an Exact Solution at the ZLB in the New Keynesian Model," NBER Working Papers 22784, National Bureau of Economic Research, Inc October 2016.

Eichenbaum, Martin and Charles L. Evans, "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates," *The Quarterly Journal of Economics*, 1995, *110* (4), 975–1009.

Engel, Charles and John H. Rogers, "The U.S. current account deficit and the expected share of world output," *Journal of Monetary Economics*, July 2006, *53* (5), 1063–1093.

Epifani, Paolo and Gino Gancia, "Openness, government size and the terms of trade," *The Review of Economic Studies*, 2009, *76* (2), 629–668.

Erceg, Christopher J. and Jesper Lindé, "Fiscal Consolidation in an Open Economy," *American Economic Review*, 2012, (102(3)), 186–191.

_ **and** _ , "Is There A Fiscal Free Lunch In A Liquidity Trap?," *Journal of the European Economic Association*, 02 2014, *12* (1), 73–107.

Evers, Michael P., "Fiscal federalism and monetary unions: A quantitative assessment," *Journal of International Economics*, 2015, *97* (1), 59–75.

Fahri, Emmanuel and Ivan Werning, "Fiscal Multipliers: Liquidity traps and currency unions," in John B. Taylor and Harald Uhlig, eds., *Handbook of Macroeconomics*, Elsevier, 2016, chapter 31, pp. 2417–2491.

Faust, Jon and John H Rogers, "Monetary policy's role in exchange rate behavior," *Journal of Monetary Economics*, 2003, *50* (7), 1403 – 1424.

Federal Ministry of Finance, "The German current account in the context of US-German trade," 2017.

Fernald, John, "A quarterly, utilization-adjusted series on total factor productivity," in "in" Federal Reserve Bank of San Francisco Working Paper 2012 -19 2014.

Forlati, Chiara, "Optimal monetary and fiscal policy in the emu: Does fiscal policy coordination matter?," *Center for Fiscal Policy, EPFL, Chair of International Finance (CFI) Working Paper*, 2009, (2009-04).

_ , "On the benefits of a monetary union: Does it pay to be bigger?," *Journal of International Economics*, 2015, *97* (2), 448–463.

Forni, Mario and Luca Gambetti, "The dynamic effects of monetary policy: A structural factor model approach," *Journal of Monetary Economics*, March 2010, *57* (2), 203–216.

Galí, Jordi, *Monetary policy, inflation, and the business cycle*, Princeton Univ. Press, 2008.

Galí, Jordi, *Monetary Policy, Inflation, and the Business Cycle: An Introduction to the New Keynesian Framework and Its Applications Second edition*, 2 ed., Princeton University Press, 2015.

Gali, Jordi and Mark Gertler, "Inflation dynamics: A structural econometric analysis," *Journal of Monetary Economics*, 1999, *44* (2), 195 – 222.

Galí, Jordi and Tommaso Monacelli, "Monetary Policy and Exchange Rate Volatility in a Small Open Economy," *Review of Economic Studies*, 07 2005, *72* (3), 707–734.

_ **and** _ , “Optimal Monetary and Fiscal Policy in a Currency Union,” Working Papers 300, IGIER (Innocenzo Gasparini Institute for Economic Research), Bocconi University 2005.

Galí, Jordi and Tommaso Monacelli, “Optimal Monetary and Fiscal Policy in a Currency Union,” *Journal of International Economics*, 2008, *76* (1), 116 – 132.

Galí, Jordi and Tommaso Monacelli, “A Note on Martin Hellwig’s “The Galí-Monacelli Model of a Small Open Economy Has No International Trade”,” 2015. mimeo.

Ghosh, Atish, “International Capital Mobility amongst the Major Industrialised Countries: Too Little or Too Much?,” *Economic Journal*, 1995, *105* (428), 107–28.

Giannoni, Marc, Christina Patterson, and Marco Del Negro, “The Forward Guidance Puzzle,” Technical Report 2016.

Gourieroux, C., A. Monfort, and E. Renault, “Indirect Inference,” *Journal of Applied Econometrics*, 1993, *8*, S85–S118.

Gourinchas, Pierre-Olivier and Aaron Tornell, “Exchange rate puzzles and distorted beliefs,” *Journal of International Economics*, December 2004, *64* (2), 303–333.

_ **and Hélène Rey**, “Chapter 10 - External Adjustment, Global Imbalances, Valuation Effects,” in Gita Gopinath, Elhanan Helpman, and Kenneth Rogoff, eds., *Handbook of International Economics*, Vol. 4 of *Handbook of International Economics*, Elsevier, 2014, pp. 585 – 645.

Gourinchas, Pierre-Olivier and Hélène Rey, “International Financial Adjustment,” *Journal of Political Economy*, 2007, *115* (4), 665–703.

Groll, Dominik and Tommaso Monacelli, “The Inherent Benefit of Monetary Unions,” CEPR Discussion Papers 11416, C.E.P.R. Discussion Papers July 2016.

Gruber, Joseph W, “A present value test of habits and the current account,” *Journal of Monetary Economics*, 2004, *51* (7), 1495–1507.

Hall, Alastair R., Atsushi Inoue, James M. Nason, and Barbara Rossi, “Information criteria for impulse response function matching estimation of DSGE models,” *Journal of Econometrics*, 2012, *170* (2), 499 – 518. Thirtieth Anniversary of Generalized Method of Moments.

Hall, Robert E., “Intertemporal Substitution in Consumption,” *Journal of Political Economy*, 1988, *96* (2), 339–357.

Herzberg, Angélique, *Sustainability of external imbalances: a critical appraisal* Springer-Gabler Research, Wiesbaden: Springer Gabler, 2015. Düsseldorf, Univ., Diss., 2013.

Huang, Chao, "An empirical study on Taiwan's current account: 1961–90," *Applied Economics*, 1993, *25* (7), 927–936.

International Monetary Fund, "External Sector Report, July 2018 : Tackling Global Imbalances amid Rising Trade Tensions," 2018.

International Monetary Fund. European Dept., "Germany: 2017 Article IV Consultation-Press Release; Staff Report; and Statement by the Executive Director for Germany," July 2017.

Jarociński, Marek and Peter Karadi, "Deconstructing monetary policy surprises: the role of information shocks," Working Paper Series 2133, European Central Bank February 2018.

Jordà, Òscar, "Estimation and inference of impulse responses by local projections," *American economic review*, 2005, *95* (1), 161–182.

Kim, Seong-Hoon, Seongman Moon, and Carlos Velasco, "Delayed Overshooting: Is It an '80s Puzzle?," *Journal of Political Economy*, 2017, *125* (5), 1570–1598.

Kim, Soyoung and Nouriel Roubini, "Exchange rate anomalies in the industrial countries: A solution with a structural VAR approach," *Journal of Monetary Economics*, 2000, *45* (3), 561 – 586.

Kirsanova, Tatiana, Mathan Satchi, David Vines, and Simon Wren-Lewis, "Optimal Fiscal Policy Rules in a Monetary Union," *Journal of Money, Credit and Banking*, October 2007, *39* (7), 1759–1784.

Kollmann, Robert, Marco Ratto, Werner Roeger, Jan in 't Veld, and Lukas Vogel, "What drives the German current account? And how does it affect other EU Member States?," *Economic Policy*, 2015, *30* (81), 47–93.

Kriwoluzky, Alexander, Gernot Müller, and Martin Wolf, "Currency Risk in Currency Unions," CEPR Discussion Papers 9635, C.E.P.R. Discussion Papers September 2013.

Krogstrup, Signe and Charles Wyplosz, "A common pool theory of supranational deficit ceilings," *European Economic Review*, 2010, *54*, 273–282.

Krugman, Paul, "The economic consequences of Herr Steinbrueck," 2008. New York Times.

Lorenzoni, Guido, "A Theory of Demand Shocks," *American Economic Review*, December 2009, *99* (5), 2050–84.

Mackowiak, Bartosz and Mirko Wiederholt, "Business Cycle Dynamics under Rational Inattention," *Review of Economic Studies*, 2015, *82* (4), 1502–1532.

Mehra, Rajnish and Edward Prescott, "The equity premium: A puzzle," *Journal of Monetary Economics*, 1985, *15* (2), 145–161.

Melosi, Leonardo, "Signalling Effects of Monetary Policy," *The Review of Economic Studies*, 2017, *84* (2), 853–884.

Mertens, Karel and Morten O. Ravn, "Understanding the aggregate effects of anticipated and unanticipated tax policy shocks," *Review of Economic Dynamics*, 2011, *14* (1), 27 – 54. Special issue: Sources of Business Cycles.

Miniane, Jacques A and Benoît Mercereau, "Challenging the Empirical Evidence From Present Value Models of the Current Account," IMF Working Papers 04/106, International Monetary Fund June 2004.

Müller, Gernot, "What drives the German current account? And how does it affect other EU Member States?," *Economic Policy*, 2015, *30* (81), 47–93.

Nakamura, Emi and Jón Steinsson, "Fiscal Stimulus in a Monetary Union: Evidence from US Regions," *American Economic Review*, 2014, *104* (3), 753–92.

_ **and Jón Steinsson**, "High-Frequency Identification of Monetary Non-Neutrality: The Information Effect*," *The Quarterly Journal of Economics*, 2018, *133* (3), 1283–1330.

_ **and** _ , "Identification in Macroeconomics," *The Journal of Economic Perspectives*, 2018, *32* (3), 59–86.

Nason, James M. and John H. Rogers, "The present-value model of the current account has been rejected: Round up the usual suspects," *Journal of International Economics*, January 2006, *68* (1), 159–187.

Newey, Whitney and Kenneth West, "A Simple, Positive Semi-definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica*, 1987, *55* (3), 703–08.

Obstfeld, Maurice and Kenneth S. Rogoff, *Foundations of International Macroeconomics*, Vol. 1 of *MIT Press Books*, The MIT Press, January 1996.

Otto, Glenn, "Testing a present-value model of the current account: Evidence from US and Canadian time series," *Journal of International Money and Finance*, October 1992, *11* (5), 414–430.

Paoli, Bianca De, "Monetary policy and welfare in a small open economy," *Journal of International Economics*, February 2009, *77* (1), 11–22.

Pappa, Evi, "Do the {ECB} and the fed really need to cooperate? Optimal monetary policy in a two-country world," *Journal of Monetary Economics*, 2004, *51* (4), 753 – 779.

Perotti, Roberto, "The Effects of Fiscal Policy in OECD countries," 2004. mimeo.

Phillips, Peter, "Impulse response and forecast error variance asymptotics in nonstationary VARs," *Journal of Econometrics*, 1998, *83* (1-2), 21–56.

Rogoff, Kenneth, "Can international monetary policy cooperation be counterproductive?," *Journal of International Economics*, 1985, *18* (3), 199 – 217.

_ , "Traded Goods Consumption Smoothing and the Random Walk Behavior of the Real Exchange Rate," 08 1992, *10.*

Rogoff, Kenneth S., *The Curse of Cash*, Princeton University Press, 2016.

Romer, Christina D. and David H. Romer, "Federal Reserve Information and the Behavior of Interest Rates," *American Economic Review*, June 2000, *90* (3), 429–457.

Romer, Christina D and David H Romer, "A new measure of monetary shocks: Derivation and implications," *American Economic Review*, 2004, *94* (4), 1055–1084.

Sachs, Jeffrey, "The Current Account and macroeconomic Adjustment in the 1970s," *Brookings Papers on Economic Activity*, 1981, *12* (1), 201–282.

_ , " The Current Account in the Macroeconomic Adjustment Process," *Scandinavian Journal of Economics*, 1982, *84* (2), 147–159.

Schmidt, Sebastian, "Optimal Monetary and Fiscal Policy with a Zero Bound on Nominal Interest Rates," *Journal of Money, Credit and Banking*, October 2013, *45* (7), 1335–1350.

_ , "Fiscal Activism and the Zero Nominal Interest Rate Bound," *Journal of Money, Credit and Banking*, June 2017, *49* (4), 695–732.

Schmitt-Grohé, Stephanie and Martın Uribe, "Closing small open economy models," *Journal of international Economics*, 2003, *61* (1), 163–185.

Schmitt-Grohé, Stephanie and Martín Uribe, "Exchange Rates and Uncovered Interest Differentials: The Role of Permanent Monetary Shocks," 2018. mimeo.

Scholl, Almuth and Harald Uhlig, "New evidence on the puzzles: Results from agnostic identification on monetary policy and exchange rates," *Journal of International Economics*, September 2008, *76* (1), 1–13.

Sheffrin, Steven M. and Wing Thye Woo, "Present value tests of an intertemporal model of the current account," *Journal of International Economics*, November 1990, *29* (3-4), 237–253.

Smets, Frank and Rafael Wouters, "Shocks and Frictions in US Business Cycles: A Bayesian DSGE Approach," *American Economic Review*, June 2007, *97* (3), 586–606.

Smith, Anthony, "Indirect Inference," 01 2008, pp. 1–6.

Soderlind, Paul, "Solution and estimation of RE macromodels with optimal policy," *European Economic Review*, April 1999, *43* (4-6), 813–823.

Steinsson, Jón, "The Dynamic Behavior of the Real Exchange Rate in Sticky Price Models," *American Economic Review*, March 2008, *98* (1), 519–533.

Stock, James H and Mark W Watson, "Identification and Estimation of Dynamic Causal Effects in Macroeconomics Using External Instruments," *The Economic Journal*, 2018, *128* (610), 917–948.

Svensson, Lars E. O., "Optimal Inflation Targets, "Conservative" Central Banks, and Linear Inflation Contracts," *The American Economic Review*, 1997, *87* (1), 98–114.

The Economist, "The German problem - Why Germany's current-account surplus is bad for the world economy," 2017. The Economist.

Turnovsky, Stephen J., "The gains from fiscal cooperation in the two-commodity real trade model," *Journal of International Economics*, August 1988, *25* (1-2), 111–127.

Werning, Ivan, "Managing a Liquidity trap: Monetary and Fiscal policy," 2012.

Woodford, Michael, "Simple Analytics of the Government Expenditure Multiplier.," *American Economic Journal: Macroeconomics*, 2011, *3* (1), 1 – 35.