Essays in International Monetary Economics

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Abstract

Chapter 2 of this thesis employs a dynamic general equilibrium with Taylor wage contracts to show that the use of strict inflation targeting as a disinflation policy may result in a slump in production and a considerable increase in macroeconomic volatility. Important determinants of the magnitude of the macroeconomic oscillations in the post-disinflation state of the economy turn out to be the size of the reduction in the inflation rate and the degree of returns to labor in the production function. In the special case of constant returns, the oscillations are large and permanent.

Chapter 3 of this thesis extends the above analysis to an open-economy setting demonstrating that the exchange rate can act as a stabilizer by effectively relieving wages from part of the burden of reducing the inflation rate. The more the economy is open, the smaller the magnitude of the macroeconomic oscillations will be after the disinflation policy is applied. The policy is shown to be infeasible for all practical purposes in a closed economy with constant returns to scale when the full nonlinear model is considered.

Chapter 4 of this thesis employs a Markov switching framework to allow for an interesting alternative characterization of macroeconomic news effects on the foreign exchange market. The chapter finds strong evidence for the presence of nonlinear regime switching between a high-volatility and a low-volatility state driven by monetary policy announcements that come as a surprise to the market. It also uncovers significant market positioning prior to the announcements, indicating a limiting of risk exposure by market participants who are unsure about the precise outcome of the policy decisions.

Chapter 5 of this thesis investigates the impact of monetary shocks on the direction and the composition of international capital flows. It identifies monetary policy shocks in a structural VAR via the pure sign restrictions approach. There are two key findings. First, a US monetary easing causes net capital inflows and a worsening of the US trade balance. Second, monetary policy shocks induce a negative conditional correlation between capital flows in bonds and equity securities. Intriguingly, they cause a negative conditional correlation between equity flows and equity returns but a positive conditional correlation between bond flows and bond returns.

Chapter 1

Introduction

A growing number of both theoretical and applied economists have devoted their time to the study of the conduct of monetary policy in recent years. Two developments stand out as factors increasing this interest.

First and foremost, progress in macroeconomic theory has been remarkable, with a new generation of quantitative models developed under the New Keynesian paradigm that can be used to explicitly study the impact of stabilization policies on economic welfare. Perhaps most importantly, these models can claim sufficient realism to be of interest to policymakers around the globe (Woodford, 2006). The origin of this influential literature lies in the synthesis of two independent theories of macroeconomic modelling. In particular, it is grounded in the idea to enrich real business cycle models of dynamic individual optimization (Prescott, 1986) with Keynesian concepts such as nominal price rigidity and the inefficiency of aggregate fluctuations. The latter had previously been developed in static models generating ultimately qualitative rather than quantitative implications (Mankiw and Romer, 1991). The development of tightly structured macroeconomic models based on explicit theoretical microfoundations and capable of successfully capturing macroeconomic time series data (Christiano, Eichenbaum and Evans, 2005; Smets and Wouters, 2003) have led to a comeback of the quantitative assessment of the optimal conduct of monetary policy (Gali, 2007).

The second reason for the popularity of the field is progress in the profession of central banking itself. A growing number of central banks have come to organize monetary policy around an explicit set of objectives. One way of specifying such objectives is via an explicit inflation forecast targeting rule as, for example, followed by the Reserve Bank of New Zealand and the Bank of England. Increased efforts to communicate policy goals to the general public also emphasize the importance of a clear framework to guide policy decisions (Woodford, 2006). In the light of these facts, it has become increasingly common to use structural macroeconomic models in policymaking institutions around the globe. These developments not only point to the success of the newly developed quantitative models but also illustrate how far the profession has come from initial efforts to organize monetary policy on an almost entirely informal basis.

The optimality of price stability as an objective for monetary policy has been at the core of this renewed interest in the study of its conduct. As demonstrated in Woodford (2003) among others, the presence of inflation not only signals an inefficient level of economic activity but also incurs a more direct cost by leading to an inefficient allocation of resources and suboptimal quantities of goods produced and consumed. From a modelling perspective, this result holds in a broad class of sticky price dynamic general equilibrium (DSGE) models (Christiano, Eichenbaum and Evans, 2005; Gali and Gertler, 1999; Woodford, 1999). At the same time, it is important to notice that there are good arguments in favour of stabilizing inflation at a strictly positive rate (e.g. in order to avoid the zero lower bound on nominal interest rates) and at the medium-term horizon (Gali, 2008). For example, Svensson (2000) shows that strict inflation targeting, a policy tolerating only minor deviations from target, leads to substantially more output volatility than a more flexible rule that targets inflation at a longer horizon.

Chapter 2 of this study somewhat reinforces this result by showing that the sub-optimality of strict inflation targeting may be even more severe when the policy is adopted as a disinflation policy, i.e. if it is applied to reduce the prevailing inflation rate to a lower level.[1] The chapter is motivated by the idea that inflation targeting differs from other disinflation policies in important respects. In particular, a strict interpretation of an inflation target allows the policymaker to tolerate only minor deviations from target. But adjusting the policy instrument such that the inflation rate is reduced to a new target and defending this target rigorously must preserve inflationary distortions such as wage and price differentials to an exceptional degree (see also Yun, 2005). The chapter employs a dynamic general equilibrium with Taylor wage contracts to show that the use of strict inflation targeting as a

[1]Rotemberg and Woodford (1997) show that in a general class of sticky price dynamic general equilibrium (DSGE) models, consumer welfare can be well approximated by a quadratic loss function in inflation and real activity. It can be shown that welfare losses are then proportional to a discounted sum of squared deviations of the current inflation rate from a moving average of recent past inflation rates, rather than deviations from zero (Sheedy, 2005). Giannoni and Woodford (2005) conclude that inflation should not be reduced too abruptly if it has been allowed to exceed its optimal long-run level.

disinflation policy may result in a slump in output and a considerable increase in macroeconomic volatility. Important determinants of the magnitude of the macroeconomic oscillations in the post-disinflation state turn out to be the size of the reduction in the inflation rate and the degree of returns to labour. In Chapter 3, the analysis is extended to an open-economy setting demonstrating that the exchange rate can act as a stabilizer by effectively relieving wages from part of the burden of reducing the inflation rate. The more the economy is open, the smaller the magnitude of the macroeconomic oscillations will be after the disinflation policy is applied.

A key merit of an inflation targeting regime under price stability is that it is believed to successfully anchor private sector expectations. Dating back to the rational expectations revolution, it is a well-accepted belief that managing private sector expectations successfully is crucial for the optimal functioning of the monetary transmission mechanism. From a macro modeling perspective, the reason is that current values of aggregate output and inflation depend not only on the central bank's current choice of the short-term interest rate, but also on the anticipated future path of this instrument (see Woodford, 2003). The practical implication is that the central bank's ability to manage private sector expectations about its future policy settings has important consequences for its overall effectiveness. Partially on the grounds of this understanding, an extensive literature developed that investigates the role of central bank communication in managing expectations (see Blinder et al (2008) for an excellent survey). The reaction of asset prices to monetary policy announcements and communication can not only be an important indicator of success in this context but can also allow uncovering important insights about the microstructure of the respective financial markets (see Sarno and Taylor, 2003).

Chapter 4 of this thesis contributes to this literature by employing a Markov switching framework in order to allow for an alternative characterization of macroeconomic news effects on the foreign exchange market. The underlying hypothesis for the choice of the model is that monetary policy announcements do not simply affect the market as shocks to an otherwise continuous process. On the contrary, news effects may change the entire data generating process underlying a market's dynamics. An econometric specification allowing for regime switches therefore appears appropriate. Indeed, one particular benefit of applying such a model is that it facilitates a plausible interpretation of observed nonlinearities. In this chapter, we find strong evidence for nonlinear regime switching between a high-volatility "informed trading" state and a low-volatility "liquidity trading" state driven by monetary policy announcements that come as a surprise to the market. We also uncover significant market positioning prior to the announcement, indicating a limiting of risk exposure by market participants.

In Chapter 5, this thesis continues to analyze the impact of monetary policy on financial markets but returns to the macro perspective. The chapter investigates the impact of monetary shocks on the direction and the composition of international capital flows. It employs a standard structural VAR specification to identify monetary policy shocks, relying on sign restrictions imposed on the impulse response functions of a few macroeconomic variables, following closely Canova and De Nicolo (2002) and Uhlig (2005). The empirical analysis yields two key findings. First, US monetary policy shocks exert a statistically and economically meaningful effect on US capital flows and the trade balance. An exogenous easing of US monetary policy by 100 basis points induces net capital inflows and a worsening of the US trade balance of around 1% of GDP after 8 quarters. The second main result focuses on the effect of monetary policy shocks on the composition of US capital flows. Intriguingly, it is found that an exogenous US monetary policy easing causes net inflows in debt securities, foreign direct investment (FDI) and other investment, while inducing net outflows in portfolio equities from the United States. Monetary policy shocks thus entail a conditional negative correlation between flows in portfolio equity and debt. A key for understanding this conditional correlation is the effect of monetary policy shocks on asset prices. While a monetary policy easing implies a decrease in short-term (and long-term) interest rates, it also causes the above mentioned increase in relative equity returns. Overall, our evidence suggests that monetary policy shocks induce negative conditional correlations between flows in bonds and equity securities. Moreover, they cause a negative conditional correlation between equity flows and equity returns and a positive conditional correlation between bond flows and bond returns.

Chapter 2

Strict Inflation Targeting as a Means of Achieving Disinflation: A Basic Analysis

2.1 Introduction

Inflation targeting has become increasingly popular as a monetary policy regime. A distinctive feature of many emerging market economies among the inflation targeters regards the level of the inflation rate exhibited at the time of adoption of the policy. Industrial countries have adopted inflation targeting at inflation rates at least broadly consistent with price stability. In contrast, emerging market economies such as Mexico, Hungary or Poland exhibited initial inflation rates of around 10 percent or more.[1] This study

[1] See Landerretche (2001), Schmidt-Hebbel and Tapia (2002), Levin et al (2004), Roger and Stone (2005) and Batini et al (2006)

is concerned with the macroeconomic effects of the explicit use of inflation targeting as a disinflation regime. The argument I make is that a strict inflation targeting policy employed for the purpose of disinflation may preserve inflationary distortions to an unusually large degree. In the model I employ, these excess distortions not only cause slumps in real activity, but may additionally increase macroeconomic volatility.

An extensive literature has analyzed inflation targeting with regard to its properties as a monetary policy regime under price stability. Its advantages have been documented by, among others, Bernanke and Mishkin (1997), Svensson (1997) and Svensson and Woodford (2003). In an early contribution, Svensson (2000) differentiates between flexible and strict inflation targeting and shows that the former creates substantially less output variability than a strict interpretation of the policy, as it effectively targets inflation at a longer horizon. The present paper somewhat reinforces this result by showing that the sub-optimality of strict inflation targeting may be even more severe in the context of a disinflation episode. Yun (2005) applies a similar line of reasoning as I do in this paper. He shows that the zero inflation optimality result (Woodford, 1999; Gali, 2000) must be refined in the presence of initial price dispersion.[2] The reason is that the pre-existing price dispersion adversely affects real activity in the economy and converges

[2]Rotemberg and Woodford (1997) show that in a general class of sticky price dynamic general equilibrium (DSGE) models, consumer welfare can be well approximated by a quadratic loss function in inflation and real activity. It can be shown that welfare losses are then proportional to a discounted sum of squared deviations of the current inflation rate from a moving average of recent past inflation rates, rather than deviations from zero (Sheedy, 2005). Giannoni and Woodford (2005) conclude that inflation should not be reduced too abruptly if it has been allowed to exceed its optimal long-run level.

faster under alternative policies.[3]

The literature on the real effects of disinflations has concentrated on explaining stylized facts regarding the differential real effects of money- and exchange-rate-based disinflations. Among others, Ball (1994) and Ascari and Rankin (2002) consider explanations for the finding that money-based disinflations typically cause slumps in output on impact. Exchange-rate-based policies, on the other hand, are frequently characterized by initial booms in real activity. Calvo and Vegh (1994) replicate this empirical finding by assuming that a collapse of the disinflation policy is rationally anticipated. Fender and Rankin (2007) explain the boom with an element of preannouncement of the policy under a standard type of exchange rate peg.[4] The long run impact of disinflation policies in the framework of the New Keynesian model is discussed in Blanchard and Gali (2007) and Ascari and Merkl (2007). The present study is motivated by the idea that inflation targeting differs from other disinflation policies in important respects. In particular, a strict interpretation of an inflation target allows the policymaker to tolerate only minor deviations from target. But adjusting the policy instrument such that the inflation rate is reduced to a new target and defending this target rigorously must preserve inflationary distortions such as wage and price differentials to an exceptional degree.

In this chapter, I use a Dynamic General Equilibrium Model with wage staggering of the type suggested by Taylor (1979a) to consider a rather ex-

[3]The stabilization policy examined in Yun (2005) does not increase macroeconomic volatility. The reason is that the author uses Calvo (1983) price contracts. I elaborate on this below.

[4]Kolver-Hernandez (2007) achieves a similar result by introducing elements of state-dependent pricing into the model economy.

treme case of a disinflation exercise: an immediate and permanent reduction in the rate of CPI inflation to a newly set target. I interpret this policy as strict inflation targeting during a disinflation episode. The central bank sets the path of money supply such that the newly set inflation target is attained immediately and sustained throughout future periods. The particular nature of the policy requires me to solve the model in a rather unconventional way. I first impose the result of the disinflation policy, a reduction to a lower rate of CPI inflation, and then solve for the policy itself, i.e. the path that money supply has to follow in order to sustain the new inflation target throughout the future.

I begin by considering a closed economy version of the model. As briefly mentioned above, I find that the disinflation policy I consider not only creates a slump in output on impact, but can additionally generate oscillatory behavior in both nominal and real variables along their post-disinflation adjustment paths. The reason is that the immediate reduction in price inflation requires the real wage to fluctuate for some periods before it gradually converges to its new steady state. The oscillations can be permanent when the returns to labor in the production function are constant. From a modeling perspective, the presence of oscillations along the adjustment path is a result that would not obtain in a model with a Calvo (1983) type staggering structure. I show that this is because price setters in such a model would always set prices equal to the prevailing price level. In the framework of model used in this study, I distinguish two cases: one in which the economy is indeed characterized by oscillatory behavior along its post-disinflation path and one in which it reacts similarly to a conventional disinflation policy. Strikingly, the size of the initial slowdown in real activity is strongly positively related to the presence of oscillations along the adjustment path.

The presence and the magnitude of oscillations along the economy's adjustment path naturally depend on the degree to which wages need to fluctuate in order to keep the inflation rate at the newly set target. Important determinants are thus the desired size of the reduction in the inflation rate as well as the returns to labor in the production function. Greater returns to scale generate oscillations as they imply that wages are tied more closely to the behavior of prices. In particular, I find that there are no oscillations at all along the post-disinflation path of the economy if the returns to labor are sufficiently low. At the other extreme, in the case of constant returns to scale, the oscillations are large and permanent. The latter is thus the only case in which the economy does not gradually converge to a new steady state.

The rest of the chapter is organized as follows. Section 2.2 outlines the structure of the model. Section 2.3 examines the implications of the disinflation policy and highlights the degree of returns to scale as a decisive factor in determining the magnitude of the oscillations in the post-disinflation state. Section 2.4 concludes.

2.2 The Model Economy

In this section, I employ a closed economy Dynamic General Equilibrium Model with imperfect competition in the labor market and nominal wage rigidities of the type proposed by Taylor (1979). The structure of the model is kept simple which allows us to illustrate our main points in a clean way and to derive crucial results analytically. I omit derivations where they are standard in the literature.

The economy is inhabited by a continuum of households $j\epsilon[0,1]$ and firms. The supply side of the economy produces a single consumption good using a technology in which labor is the only variable factor of production

$$Y_t = N_t^{\sigma} \tag{2.1}$$

where Y_t is output at time t, σ is the degree of returns to labor and $0 < \sigma \leq 1$. A typical firm demands a continuum of labor types $j\epsilon[0,1]$ to minimize the cost of achieving a particular composite labor input N_t, given by $N_t = [\int_0^1 L_{jt}^{(\varepsilon-1)/\varepsilon} dj]^{\varepsilon/(\varepsilon-1)}$. L_{jt} is the quantity of labor that household j supplies to the firm and $\varepsilon > 1$ is the elasticity of technical substitution across labor types. Solving the cost minimization problem yields the standard conditional demand for labor function

$$L_{jt} = N_t(\frac{W_t}{W_{jt}})^{\varepsilon} \tag{2.2}$$

where W_t is the wage index given by $W_t = [\int_0^1 W_{jt}^{1-\varepsilon} dj]^{1/(1-\varepsilon)}$. Goods markets are perfectly competitive. The firm's profit maximization problem thus yields the supply function

$$Y_t = N_t(\frac{W_t}{\sigma P_t})^{\sigma/(\sigma-1)} \tag{2.3}$$

I now move to the demand side of the economy. Household j supplies labor skill $j\epsilon[0,1]$ and sets its own wage W_{jt}. I assume that the economy consists of two sectors of households. Sector A comprises labor types $[0, 0.5)$ and sector B comprises labor types $[0.5, 1]$. Although households are monopolistic suppliers of their individual type of labor input, they are price takers in all other markets. I assume that they are completely symmetric in terms of their preference structure which implies that consumption must be equal across households in a given sector at any point in time. I also assume the existence of complete domestic asset markets. This implies that households can insure against any type of initial shock that might affect the two sectors differently due to the staggering structure to be defined below. Hence, $C_{jt} = C_{kt}$ must hold for any two households j and k in sectors A and B. Finally, I define aggregate nominal consumption as $S_t = P_t C_t$ where P_t is the price of one unit of the composite consumption good.

Wages are set by each individual household subject to a staggering structure of the type proposed by Taylor (1979). In particular, I assume that households in sector A (B) set their wage in even (odd) periods and keep it fixed for the subsequent period. The wage newly set in period t is denoted by X_t independently of the sector in which it is set. Households are utility maximizers. A representative household j in sector A derives utility from consumption, liquidity holdings and leisure and maximizes her discounted lifetime utility U_j by choosing a pattern for personal consumption C_{jt}, bond holdings B_{jt}, wages W_{jt} and labor effort N_{jt} subject to a series of budget constraints, the conditional demand for labor and the wage setting constraint:

$$U_j = \sum_{t=0}^{\infty} \beta^t [\delta ln C_{jt} + (1-\delta) ln(M_{jt}/P_t) - \eta L_{jt}^{\zeta}] \tag{2.4}$$

subject to

$$M_{jt-1} + I_{t-1}B_{jt-1} + W_{jt}L_{jt} + \Pi_t + G_t = P_tC_{jt} + M_{jt} + B_{jt} \tag{2.5}$$

$$L_{jt} = N_t(\frac{W_t}{W_{jt}})^{\varepsilon} \tag{2.6}$$

$$W_{jt} = W_{jt+1} = X_t, \;\; t = 0, 2, 4, ... \tag{2.7}$$

where only the first two constraints must hold in even and odd periods $t = 0, 1, 2, 3, ...$ and where $\beta < 1$, $\zeta \geq 1$, I_t is the domestic gross interest rate, M_t denotes money supply in period t, G_t is a lump-sum subsidy to households and Π_t denotes a share in firms' profits that is equal across households.[5] The optimization problem for household k in sector B is exactly equivalent except that the wage setting constraint holds in odd instead of even periods. The first order conditions of this optimization problem are given by the consumption Euler equation, the money demand optimality condition and the optimal wage setting condition

$$C_{jt+1} = \beta[I_t\frac{P_t}{P_{t+1}}]C_{jt} \tag{2.8}$$

$$\frac{M_{jt}}{P_t} = C_{jt}\frac{1-\delta}{\delta}\frac{I_t}{I_t - 1} \tag{2.9}$$

[5] A no-Ponzi-game condition ensures that individuals cannot borrow infinitely by repaying debt with further debt.

$$X_t = [\frac{\varepsilon}{\varepsilon - 1}\frac{\eta\zeta}{\delta}\frac{L_{jt}^{\zeta} + \beta L_{jt+1}^{\zeta}}{\frac{L_{jt}}{P_t C_{jt}} + \beta\frac{L_{jt+1}}{P_{t+1}C_{jt+1}}}] \quad (2.10)$$

The economy further comprises a government that controls money supply. Its budget constraint is given by

$$G_t = M_t - M_{t-1} \quad (2.11)$$

The competitive equilibrium in the model economy is the sequence of prices $[X_{jt}, P_t]_{t=0}^{\infty}$ and allocations $[Y_t, C_{jt}, N_{jt}, L_{jt}, B_{jt}, M_{jt}]_{t=0}^{\infty}$ such that firms maximize profits, agents maximize utility and all markets clear.[6] I aggregate the equilibrium conditions across individuals and take a log-linear approximation around a reference steady state in which inflation is zero. Notice that in what follows I will distinguish between the reference zero inflation steady state (ZISS) around which the equilibrium conditions are linearized and a constant inflation steady state (CISS) in which the economy finds itself one period before the disinflation policy is applied. As the economy is closed, bond holdings are zero in both states.[7] The linearized equilibrium conditions are presented in Appendix A.1.1.[8] In the following, lower-case symbols denote log deviations of variables from their reference steady state values, i.e. $v_t = log\frac{V_t}{V_R}$.

[6]Due to the staggering structure in wages the labor market does not clear in the Walrasian sense.

[7]While this must necessarily be the case in a closed economy setting, I will impose it upon the open economy by assumption in the next section

[8]The interested reader may refer to Fender and Rankin (2008) for the linearization of the wage setting condition.

2.3 Strict Inflation Targeting as a Disinflation Policy

I now proceed to assess the macroeconomic effects of a disinflation policy that uses strict inflation targeting to reduce the rate of CPI inflation from a positive to a non-negative value. I define strict inflation targeting following Svensson (2000). The definition implies that the policymaker only cares about stabilizing inflation at a given target. I assume that she has perfect control over the inflation rate and adjusts her policy instrument such that inflation is kept at target at any point in time. The economy is initially in a constant inflation steady state (CISS) in which real variables are constant and all nominal variables grow at some constant inflation rate μ_I. The policymaker decides to reduce inflation in period $t = 0$ from its initial rate μ_I to the lower but nonnegative value μ_D. This policy change is unexpected and credible. The policymaker takes action by announcing μ_D as the new target and adjusts the path of money supply such that the new target is attained immediately and sustained throughout future periods.[9]

There is no doubt that this disinflation policy is more rigid than policies applied in the real world. However, I present this extreme example as a benchmark case that allows for an analytical characterization of possible adverse effects of applying a too rigid inflation targeting policy for the purpose of disinflation. The definition of the policy requires us to solve the model in a rather unconventional way. I first impose the result of the disinflation policy,

[9]There is a unique path of money supply, the exchange rate and the nominal interest rate that achieves this outcome. It is therefore not of importance, whether we think of the central bank as using one or the other as its policy instrument.

a reduction in the rate of inflation from μ_I to μ_D, and then solve for the path that the money supply needs to follow in order to achieve and sustain the newly set inflation target. Finally, I assess the macroeconomic effects of the disinflation policy.

2.3.1 The Initial Constant Inflation Steady State

Before I can investigate the macroeconomic impact of the disinflation policy, I need to solve for the equilibrium of the model in the initial CISS. I use the fact that the new wage x_t is homogeneous of degree one in nominal variables and, when normalized by money supply such that $v_t = x_t - m_t$, is constant over time at

$$v = -\frac{1}{2\gamma}\big[-\gamma\frac{1+\beta}{1-\beta} + \frac{1-\beta}{1+\beta}\big]\mu_I \tag{2.12}$$

with $\gamma = \frac{\zeta}{1+\varepsilon(\zeta-1)}$. The CISS value of nontradables output is then given by

$$y = \frac{\sigma(1-\beta)}{2(1+\beta)\gamma}\mu_I \tag{2.13}$$

We observe that nontradables output is affected positively by the inflation rate in the CISS. The effect is due to wage setters discounting future utility. I will further discuss this effect in the next chapter in the context of the full nonlinear model. In the CISS, all nominal variables grow at the rate μ_I. Defining m_{-1} as the level of money supply one period before the disinflation policy is applied, I can write the wage index in the same period as

$$w_{-1} = -\frac{1}{2}\left[\frac{-2\beta}{1-\beta} + \frac{1-\beta}{(1+\beta)\gamma}\right]\mu_I + m_{-1} \tag{2.14}$$

Equations (2.13) and (2.14) constitute the CISS solution of the model one period before disinflation.

The next step in the analysis is to understand, how changes in the rate of price inflation affect the rest of the economy and in particular the staggered variable, wages. The supply function (A.5) allows me to write the price index as a function of output and the wage index

$$p_t = \frac{1-\sigma}{\sigma} y_t + w_t \tag{2.15}$$

The equation illustrates that a given change in the price level is accommodated partly by a reaction of wages and partly by firms adjusting production. The wage level's share in facilitating the price change depends positively on the degree of returns to labor. The reason is that a larger σ implies a less strongly upward sloping marginal cost curve and thus a weaker response of prices to changes in the level of production in the economy. When $\sigma = 1$, the marginal cost curve is horizontal, leaving no role for output in the determination of prices. This implies that the wage level must move one for one with the price index. In the context of our disinflation policy the latter case corresponds to a situation in which a reduction in the rate of price inflation translates one for one into a reduction in the growth rate of the wage index. The subsequent section will show that this property allows for a purely analytic characterization of the disinflation policy's impact on the economy.

2.3.2 The Special Case: Constant Returns to Scale

I begin by solving for the impact of the disinflation policy in the special case of constant returns to scale. I do so because the distinctive implications of this assumption allow for a good understanding of the channels through which the policy affects the economy.

As a starting point, it is important to notice that there is a crucial implication arising from the assumption of constant returns to scale in the production function. In particular, equation (A.5) shows that this is the one and only case for which it is not possible to solve for output when the paths of both the wage index and the price level are known. The reason is that the aggregate supply curve, i.e. the relationship between output and the price level, is horizontal for a given wage. This implies that conventional solution techniques cannot be applied to determine the full post-disinflation solution of the model, a point that will perhaps become more obvious when I examine the general case of $\sigma < 1$ in the next section.

There is, however, another anomaly about the case of constant returns to scale. In the previous subsection I have shown that $\sigma = 1$ implies that the wage index moves one for one with the price index and a policy of reducing CPI inflation effectively becomes one of reducing the growth rate of the wage index. Using this fact, it turns out to be possible to directly solve for the post-disinflation path of the real side of the economy by distinguishing the economy's law of motion in even periods from the law of motion in odd periods. To see this, remember that our disinflation policy reduces CPI inflation from its initial rate μ_I to a lower but non-negative rate μ_D. Given that the wage index moves one for one with prices, it is easy to see what such a policy implies for the behavior of sectoral wages. I take the wage index and its one period lag. I evaluate the difference at $t = 0$. This yields

$$w_0 - w_{-1} = \frac{1}{2}(x_0 - x_{-1}) + \frac{1}{2}(x_{-1} - x_{-2}) \tag{2.16}$$

I have assumed that the economy is in a CISS up until period $t = 0$. It follows that the second term on the RHS is predetermined and equal to $\frac{1}{2}\mu_I$. Reducing inflation to μ_D in period $t = 0$ then implies that the LHS of (2.16) is equal to μ_D. Hence,

$$x_0 - x_{-1} = 2\mu_D - \mu_I \tag{2.17}$$

Following the equivalent procedure for the subsequent periods, we have

$$x_t - x_{t-1} = 2\mu_D - \mu_I \tag{2.18}$$

for all $t = 0, 2, 4, \ldots, \infty$, and

$$x_t - x_{t-1} = \mu_I \tag{2.19}$$

for all $t = 1, 3, 5, \ldots, \infty$. In other words, wages grow at the constant rate $2\mu_D$ from one even period to the other and from one odd period to the other. But, since $\mu_D < \mu_I$ the new wage grows faster from even to odd periods than from odd to even periods. Three cases are possible. First, if $2\mu_D < \mu_I$, the new wage set in even periods is lower than the wage set in the previous odd period. Wages in sector A are thus smaller than wages in sector B throughout

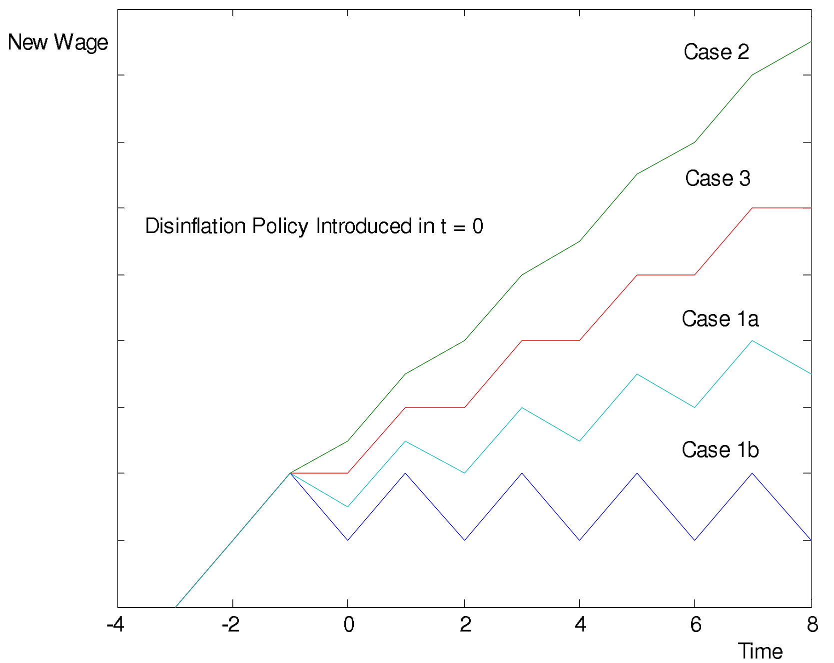


Figure 2.1: Post-Disinflation Path of the New Wage

the post-disinflation state. Second, if $2\mu_D > \mu_I$, the new wage set in even periods is higher than the wage set in the previous odd period. Third, if $2\mu_D = \mu_I$, the new wage set in even periods is equal to the prevailing one from the previous odd period.

Figure 2.1 illustrates these three cases. The special case of $\mu_D = 0$ is represented by Case 1b. It represents a complete disinflation, i.e. a reduction in the rate of wage inflation from μ_I to zero. In particular, $\mu_D = 0$ implies that

$$x_{-2} = x_0 = x_2 = ... = w_{-1} - \frac{1}{2}\mu_I \tag{2.20}$$

$$x_{-1} = x_1 = x_3 = ... = w_{-1} + \frac{1}{2}\mu_I \quad (2.21)$$

That is to say that the disinflation policy produces a wage gap between the two sectors of households that is constant and permanent. The intuition is the following: I imposed zero wage inflation from period zero onwards. The wage set in period $t = -2$ is smaller than the one set in period $t = -1$. This implies that the wage index requires downward pressure by the new wage set in period zero to ensure zero wage inflation. Due to the fact that the wage index is an equally weighted average of the two sectoral wages, the necessary amount of downward pressure is achieved if the wage in period $t = 0$ is set equal to wage set in period $t = -2$. The opposite reasoning applies in the subsequent period.

Let us get back to the more general case of an inflation reduction from a constant value μ_I to a lower positive value μ_D. Having determined the wage setting behavior of households after the disinflation policy is introduced, I can now derive the entire post disinflation state of the economy. We substitute for both the wage index and employment in the wage setting condition (A.6) and obtain

$$x_t = \frac{1}{1+\beta}[s_t \frac{2\gamma}{1+\gamma} + x_{t-1}\frac{1-\gamma}{1+\gamma}] + \frac{\beta}{1+\beta}[s_{t+1}\frac{2\gamma}{1+\gamma} + x_{t+1}\frac{1-\gamma}{1+\gamma}] \quad (2.22)$$

We already know how wages behave in the post-disinflation state of the economy. In principle, we can simply impose their paths on the law of motion of the economy. However, it is important to notice that the dynamic path of sectoral wages differs between even and odd periods. This implies that we have to look at each period separately. Using the wage setting condition (A.6) and substituting from equations (2.20) and (2.21), we have that

$$s_t + \frac{1}{\beta}s_{t-1} = \frac{1+\beta}{\beta}(w_{-1} + \frac{1}{2}\mu_I + t\mu_D) - \frac{1-\gamma}{2\gamma}(2\mu_D - \frac{1+\beta}{\beta}\mu_I) \quad (2.23)$$

$$s_t + \frac{1}{\beta}s_{t-1} = \frac{1+\beta}{\beta}(w_{-1} - \frac{1}{2}\mu_I + (t+1)\mu_D) - \frac{1-\gamma}{2\gamma}(\frac{1+\beta}{\beta}\mu_I - \frac{1}{\beta}2\mu_D) \quad (2.24)$$

in odd and even periods respectively. These relationships hold from period zero onwards. A similar relationship between s_0 and s_{-1} would be convenient as it would uniquely determine the post disinflation path of s_t. However, the equivalent procedure would give us a relationship between s_{-1} and the expectation of s_0 as of period $t = -1$. And this expectation is false as soon as the disinflation policy is introduced. The reason is that, in period $t = -1$, the wage setter does, by assumption, not know that the disinflation policy will be applied. Nominal consumption in the subsequent period then turns out to be different from her expectation as of period $t = -1$. But we can get around this problem by appealing to stability reasoning. From (2.23) and (2.24), we have that in odd and even periods respectively

$$s_{t+2} - \frac{1}{\beta^2} s_t = \mu_D \frac{\beta^2 - 1}{\beta^2} t + \frac{\beta^2 - 1}{\beta^2} w_{-1} - \mu_I (\frac{1+\beta}{\beta})^2 \frac{1}{2\gamma} + \mu_D (3 + \frac{2\beta - \gamma}{\beta^2 \gamma}) \tag{2.25}$$

$$s_{t+2} - \frac{1}{\beta^2} s_t = \mu_D \frac{\beta^2 - 1}{\beta^2} t + \frac{\beta^2 - 1}{\beta^2} w_{-1} + \mu_I (\frac{1+\beta}{\beta})^2 \frac{1}{2\gamma} + \mu_D (3 - \frac{\gamma + 1 + \beta^2}{\beta^2 \gamma}) \tag{2.26}$$

Equations (2.25) and (2.26) are the basic laws of motion of the economy in the post disinflation state. In order to arrive at a stationary solution, I deflate both difference equations by the wage level w_t. I create a new variable a_t defined as the ratio $a_t = s_t - w_t$. The resulting first order difference equations in a_t take the form

$$a_{t+2} = \frac{1}{\beta^2} a_t + constant \tag{2.27}$$

Since $\beta < 1$, the Eigenvalue of these equations is greater than one and unstable. This implies that there is a unique non divergent solution for each of these processes as Blanchard and Kahn (1980) show. This non divergent solution obtains when a_t is equal across odd and across even periods. The economy instantly jumps to its post disinflation state and remains there. Using the goods supply and demand functions, I find

$$y_{odd} = \frac{(1+\beta)}{2(1-\beta)\gamma}\mu_I - \frac{2\beta}{(1-\beta^2)\gamma}\mu_D \tag{2.28}$$

$$y_{even} = -\frac{(1+\beta)}{2(1-\beta)\gamma}\mu_I + \frac{(1+\beta^2)}{(1-\beta^2)\gamma}\mu_D \tag{2.29}$$

As is evident from equations (2.28) and (2.29), the disinflation policy results in output oscillations between even and odd periods. The amplitude of these oscillations is constant. Moreover, we can observe that the extent to which nominal and real variables fluctuate depends positively on the initial inflation rate and negatively on the post-disinflation rate. The reason is that the inflation rate determines the degree of wage dispersion present in the initial CISS. This suggests that a more gradual path of disinflation may attenuate the increase in output volatility resulting from the disinflation policy. The intuition for these findings becomes clear when the post disinflation path of money supply is derived. I have explained above that the policy instrument is endogenous in our solution procedure although it is exogenous in the interpretation of the policy experiment. Notice first that nominal consumption can be expressed as

$$s_t = y_t + w_t \tag{2.30}$$

Given the path of output documented in (2.28) and (2.29), and given that the wage index grows at the rate μ_D from period to period, we observe that post disinflation nominal consumption also grows at a lower rate in even periods than it does in odd periods. The post-disinflation path of money supply is given as a function of the path of nominal consumption

$$m_t = \frac{\beta}{1-\beta}(\frac{1}{\beta}s_t - s_{t+1}) \tag{2.31}$$

It is easy to see that money supply must follow a pattern that is qualitatively similar to the path of wages and output. The intuition is the following: the disinflation policy effectively imposes a path for the new wage that requires agents to set low wages in even periods and high wages in odd periods. In order to motivate this wage setting behavior, money supply must be set low in even and high in odd periods. This also explains the oscillatory behavior of output. Finally, it is straightforward to show that average output in the post-disinflation state is lower than output in the initial CISS.[10] This is to say that the disinflation policy creates a slump in production.[11]

2.3.3 The General Case: Decreasing Returns to Scale

I now abstract from the assumption of constant returns to scale and return to the more general case of $\sigma < 1$. The analysis in the previous section has shown that the disinflation policy can result in oscillations in sectoral real wages that are permanent and constant over time. In this section, I show that this result is not robust to relaxing the assumption of $\sigma = 1$. In particular, I find that the economy ultimately converges to a new steady state in this more

[10] But output is greater than or equal to that in the ZISS. The reason is that the discounting effect of inflation on output is not only present in the initial CISS but also, albeit attenuated, in the post-disinflation state.

[11] It is typically found that money based disinflations cause a slump in output on impact, while exchange rate based disinflations can cause a boom. Explanations for these empirical findings can be found in Calvo and Vegh (1994), Rebelo and Vegh (1995), Fender and Rankin (2006) or Kolver Hernandez (2007).

general case. The transition path may or may not be subject to oscillations, depending on the particular value of σ.

I investigate the effects of the disinflation policy in a different way than in the previous section. The reason is that it is now possible to derive the entire post-disinflation path of the economy by determining the response of the staggered variable, wages, to the reduction in price inflation. I again consider a reduction in $p_t - p_{t-1}$ from μ_I to μ_D in period zero. In particular, taking equation (2.22) and substituting for nominal consumption s_t using the definition (A.2) and the supply function (A.5), I can express the new wage in period t solely as a function of past and future wages as well as the price index

$$\begin{aligned} x_t = & \frac{1}{1+\beta}[\frac{2\gamma}{1+\gamma}(\frac{1}{1-\sigma}p_t - \frac{1}{2}\frac{\sigma}{1-\sigma}(x_t + x_{t-1})) + \frac{1-\gamma}{1+\gamma}x_{t-1}] \\ & + \frac{\beta}{1+\beta}[\frac{2\gamma}{1+\gamma}(\frac{1}{1-\sigma}p_{t+1} - \frac{1}{2}\frac{\sigma}{1-\sigma}(x_{t+1} + x_t)) + \frac{1-\gamma}{1+\gamma}x_{t+1}] \end{aligned} \quad (2.32)$$

In the post-disinflation state p_t and $p_{t+1} = p_t + \mu_D$ are exogenous and predetermined by definition of the disinflation policy. Therefore, the households wage setting condition can, from period $t = 0$ onwards, be expressed as

$$x_{t+1} - \frac{1+\beta}{\beta}\frac{1-\sigma+\gamma}{1-\sigma-\gamma}x_t + \frac{1}{\beta}x_{t-1} = -\frac{1+\beta}{\beta}\frac{2\gamma}{1-\sigma-\gamma}(p_t + \frac{\beta}{1+\beta}\mu_D) \quad (2.33)$$

This is the law of motion of the economy in the post-disinflation steady state. Notice that p_t is growing over time so that we need to stationarize the equation before we can solve it. I do so by defining $\phi_t = x_t - p_t$ and obtain

$$\phi_{t+1} - \frac{1+\beta}{\beta}\frac{1-\sigma+\gamma}{1-\sigma-\gamma}\phi_t + \frac{1}{\beta}\phi_{t-1} = \mu_D(\frac{1-\beta}{\beta} - \frac{2\gamma}{1-\sigma-\gamma}) \tag{2.34}$$

As I show in Appendix A.1.2, equation (2.34) is saddlepoint stable when $\sigma < 1$, i.e. one of its eigenvalues lies within and the other outside the unit circle. This result holds for any choice of parameter values within the defined limits.[12] I use the eigenvalue-eigenvector solution technique of Blanchard and Kahn (1980) to solve for the post-disinflation path of ϕ_t. As (2.34) is a scalar second order difference equation and its RHS is constant over time, its rational expectations solution is given by

$$\phi_t = \lambda_1\phi_{t-1} - \mu_D\frac{(1-\beta)/\beta - 2\gamma/(1-\alpha\sigma-\gamma)}{\lambda_2 - 1} \tag{2.35}$$

where λ_1 denotes the eigenvalue that is smaller in absolute value and λ_2 denotes the one that is bigger. Knowing both the path of the price index and the normalized new wage ϕ_t, it is straightforward to derive the entire post-disinflation path of the economy. In reference to the previous section, notice that this would not have been possible in the special case of constant returns to scale in the production function. The reason is that the aggregate supply curve in this case is horizontal for a given wage level, implying that

[12] When $\sigma = 1$, one of the two eigenvalues is equal to -1, which is consistent with the finding that there are permanent oscillations in the case of constant returns to scale.

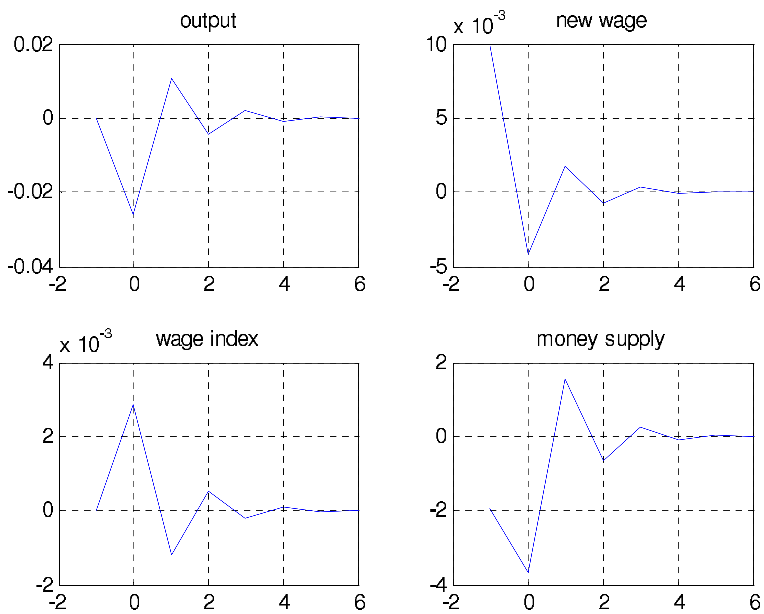


Figure 2.2: Closed Economy: Post-Disinflation Path with Oscillations

the post-disinflation path of the real economy could not have been derived from the sole knowledge of the path of prices and wages.

Whether the path of the economy is subject to oscillations as in the case of wage inflation targeting now crucially depends on the sign of the smaller eigenvalue. Appendix A.1.2 shows that the smaller eigenvalue is of negative sign, and thus induces oscillations, if and only if $1 - \sigma - \gamma < 0$ holds, where $\gamma = \frac{\zeta}{1+\varepsilon(\zeta-1)}$. This is to say that the post-disinflation path of the economy exhibits oscillations for a sufficiently large σ and sufficiently small ζ and ε.

I study the impact of the disinflation policy in both situations using the example of an initial inflation rate of two percent and a complete elimination in price inflation after the policy is applied. I first investigate the case in which $1 - \sigma - \gamma < 0$ holds.[13] Figure 2.2 illustrates the resulting path of

[13] I choose a rather extreme parametrization with $\sigma = 0.9$, $\zeta = 1.3$ and $\varepsilon = 4$.

the economy. As in the case of constant returns to scale, we observe strong oscillations in both nominal and real variables after the disinflation policy is applied. The new wage falls on impact, oscillates for a few periods, and then gradually converges to its ZISS value. Money supply and output behave similarly. Moreover, average output from period zero onwards is substantially lower than prior to the application of the policy. The policy therefore not only increases macroeconomic volatility but also creates a slump in output. In the impact period of the policy, there is a drop of more than 2 percent in real activity.

These results are somewhat similar to the previously discussed special case of constant returns to scale in the production function. However, the crucial difference is that the economy converges over time after some periods of increased volatility. The reason is related to the assumption of decreasing returns to labor. As discussed previously, firms now face upward sloping marginal cost curves, which relieves wages of part of the burden of responding to desired changes in the price level or its growth rate.[14] And the weaker is the response of wages in the impact period of the shock, the weaker will be the response in the subsequent period. Hence, the magnitude of the fluctuations in the economy decreases over time. The smaller is σ, the more rapidly this process takes place. The attenuating effect is strengthened the greater is the elasticity of substitution between labor types ε and the greater is the elasticity of the disutility of labor ζ. [15]

[14]The intuition is simple: in the presence of an upward sloping marginal cost curve, the initial fall in output relieves wages from part of the burden of responding to the change in prices.

[15]Regarding ε, the reason is that a higher substitutability of labor types implies that a given firm will employ more labor of the low wage type in period $t = 0$. This reduces

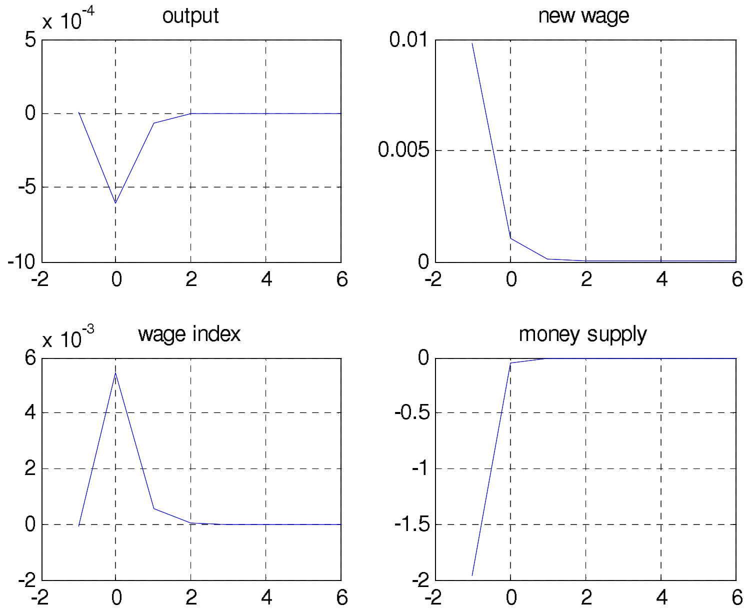


Figure 2.3: Closed Economy: Post-Disinflation Path without Oscillations

Let us now focus on the case of $1 - \sigma - \gamma > 0$ in which σ is small enough and ε and ζ are large enough such that there are no oscillations in the post-disinflation state. I choose $\sigma = 0.1$ and leave all remaining parameter values the same as in the previous case. Figure 2.3 shows that there are now indeed no oscillations in the post-disinflation path of the economy. All variables monotonically converge to their ZISS levels. In fact, the post-disinflation path of output reminds us of the one that is typically attained under a conventional disinflation policy in a model with a staggering structure in wages or prices.

To sum up, the analysis has shown that a strict inflation targeting policy employed to reduce CPI inflation will result in a slump in output on impact

the wage index and implies that the new wage set in period $t = 0$ does not have to be as low as would have been the case otherwise. A higher disutility per unit of work effort, on the other hand, scales wages upward and thus also scales the wage gap between one sector and the other.

and may create oscillations of a substantial magnitude in both nominal and real variables. The presence of oscillations in the post-disinflation path of the economy crucially depends on the degree of returns to labor in the production function. For a large enough σ the disinflation policy will result in oscillatory behavior of both real and nominal variables. These oscillations die out more slowly over time the larger is σ and are permanent in the limiting case of $\sigma = 1$. Finally, notice that the slump in output following the disinflation policy is tiny when there are no post-disinflation oscillations in the economy and substantial if there are.

2.3.4 Digression: Calvo Staggering Structure

In this section I briefly show that the post-disinflation oscillations in response to our disinflation policy would not obtain in a model with a staggering structure as proposed by Calvo (1983). This is the reason why the policy experiment of Yun (2005) does not lead to an equivalent conclusion.

Following Yun (2005) and much of the literature, let us consider the case of price inflation targeting in an economy in which prices are subject to a Calvo-type staggering structure. We can write the price index in the absence of indexation as

$$P_t^{1-\varepsilon} = (1-\alpha)P_{t,t}^{1-\varepsilon} + \alpha P_{t-1}^{1-\varepsilon} \tag{2.36}$$

where P_t is the price index and $P_{t,t}$ is the price that is chosen by the fraction $(1-\alpha)$ of firms that get to adjust their price in period t. Let us assume a policy in which price inflation is reduced to zero once and for all in some period $t = 0$. Computing the ratio of P_t and P_{t-1} and imposing $\frac{P_t}{P_{t-1}} = 1$, I find that $P_{t,t} = P_{t-1}$. Thus, newly set prices in the post-disinflation state of the economy are always equal to the prevailing price level. The reason is that firms choose prices in a forward looking manner and face the identical problem in every period. Since prices gradually converge, there is no source for oscillatory behavior in the economy. The increase in macroeconomic volatility resulting from the disinflation policy applied in our model is thus a result that does not obtain in Calvo-type staggering models.

2.4 Discussion

The present study is motivated by the idea that inflation targeting differs from other disinflation policies in important respects. In particular, a strict interpretation of an inflation target allows the policymaker to tolerate no deviations from target. But adjusting the policy instrument such that the inflation target is attained and defending the new target rigorously may preserve inflationary distortions such as wage and price differentials to an exceptional degree. The implications of these distortions might be severe for both real activity and volatility in the economy.

In order to formalize this idea, I have used a Dynamic General Equilibrium Model with wage staggering of the type suggested by Taylor (1979a) to consider a rather extreme case of a disinflation policy: an immediate and permanent reduction in the rate of CPI inflation to a newly set target. I found that the disinflation policy not only creates a slump in output on impact, but can additionally generate oscillatory behavior in both nominal and real variables along their post-disinflation adjustment path. These oscillations can be permanent when the returns to labor in the production function are constant. In this case only, the economy does not gradually converge to a new steady state. Moreover, I showed that the size of the initial slowdown in real activity and the magnitude of the oscillations are positively related.

From a modeling perspective, the analysis has shown that the presence of oscillations along the adjustment path would not occur in a model with a Calvo (1983) type staggering structure as in Yun (2005). In the framework of the particular model I employed, I found that the presence of oscillations along the post-disinflation path of the economy as well as their magnitude strongly depend on the desired size of the reduction in the inflation rate as well as the returns to labor in the production function. In particular, I illustrated that there are no oscillations at all along the post-disinflation path of the economy if the returns to labor are sufficiently low. At the other extreme, in the case of constant returns to scale, the oscillations are large and permanent. The reason is that, in the latter case, CPI inflation targeting becomes equivalent to a policy of wage inflation targeting.

The present study is in line with the analysis of Yun (2005) in that it identifies the slow convergence of prices as the major source of inefficiency resulting from an unexpected, immediate and permanent reduction in the inflation rate. Moreover, the analysis has shown that the negative consequences of strict inflation targeting identified by Svensson (2000) may be exacerbated when the policy is used as a disinflation regime. I interpret this as an explanation for the finding of Roger and Stone (2005) that target misses are particularly common for disinflating inflation targeters despite the fact that monetary authorities in these economies should be particularly eager to avoid credibility losses.[16] In sum, there are good reasons to be cautious when adopting inflation targeting as a disinflation regime. This may be particularly true for emerging market economies which have a limited experience with an independent monetary authority.

[16]The authors provide evidence that suggests that central banks rather shift up the planned trajectory for the inflation rate than tightening policy in order not to deviate from it. This behavior complies with the concept of opportunistic disinflation proposed by Clifton (1999).

Chapter 3

Strict Inflation Targeting as a Means of Achieving Disinflation: Extensions

3.1 Introduction

In this chapter, I investigate how the conclusions from the previous chapter change in an open economy setting. In addition to the returns to scale in the production function, the degree of openness of the economy is identified as a decisive factor in determining the presence and the magnitude of the oscillations in the post-disinflation state of the economy. I show that the degree of openness matters because the exchange rate acts as a stabilizer along the post-disinflation path of the economy by effectively relieving wages of part of the burden of reducing the inflation rate. The more open is the economy, the greater is the share of the burden it can successfully manage.

Furthermore, I investigate the impact of the disinflation policy on the economy in the full nonlinear model (as opposed to using a loglinear version of it as in the previous and the first part of the current chapter). Although the results are qualitatively very similar to the loglinear economy, a set of additional conclusions emerges in the special case of a closed economy with constant returns to scale in the production function. It turns out that the policymaker faces a surprisingly strict feasibility constraint that does not allow for the policy to be carried out in virtually any case of empirical relevance.

The rest of the chapter is organized as follows. In Section 3.2, I extend the model economy employed in the previous chapter to an open economy framework. Section 3.3 identifies the degree of openness as a decisive factor in determining the impact of the disinflation policy on macroeconomic volatility. Section 3.4 investigates the disinflation policy in the framework of the full nonlinear model. Section 3.5 concludes.

3.2 Opening up the Model Economy

I now assume that the economy is open in the sense that its agents trade goods and assets with a foreign country. The economy's basic structure is the same as in the previous section. In opening up the economy I follow the formulation of Fender and Rankin (2008).

I assume that there are now two output sectors, one producing tradable goods and one producing non-tradable goods. Output in the tradables sector Y_{Tt} is exogenous and normalized to one. The production function for non-tradable goods Y_{Nt} is equivalent to equation (2.1) in the closed economy specification. Markets for both types of goods are perfectly competitive. The foreign currency price of tradables is normalized to unity. Together with the assumption that the law of one price holds, this implies that $P_{Tt} = E_t$, where E_t is the nominal exchange rate, i.e. the domestic price of foreign currency. As regards financial markets, B_t is now an international bond traded between home and foreign agents. The currency of its denomination is immaterial since we assume that there are no initial outstanding bonds and there is no uncertainty after the disinflation policy is applied. A no-arbitrage condition implies interest rate parity

$$I_t = I_t^* \frac{E_{t+1}}{E_t} \tag{3.1}$$

where I_t^* is the foreign gross interest rate. The optimization problems of the individual agents are equivalent to the closed economy case except that households now consume both non-tradable and tradable goods. Their preference structure is revealed by the composite consumption index

$$C_{jt} = C_{Njt}^{\alpha} C_{Tjt}^{1-\alpha} \tag{3.2}$$

where $0 < \alpha < 1$ and C_{Njt} and C_{Tjt} denote household j's consumption of nontradables and tradables respectively. Utility from consumption is maximized subject to a given nominal spending constraint S_{jt} defined by $S_{jt} = P_{Nt}C_{Njt} + P_{Tt}C_{Tjt}$, such that

$$C_{Njt} = \alpha S_t / P_{Nt} \tag{3.3}$$

$$C_{Tjt} = (1-\alpha)S_t/P_{Tt} \tag{3.4}$$

where α is the degree of home bias in consumption. I denote $1-\alpha$ the degree of openness of the economy. The consumer price index (CPI) is then a weighted average of the price of tradables and non-tradables. In particular,

$$P_t = \frac{P_{Nt}^{\alpha} P_{Tt}^{1-\alpha}}{\alpha^{\alpha}(1-\alpha)^{1-\alpha}} \tag{3.5}$$

The trade balance T_t can be expressed as $T_t = 1 - C_{Tt}$. In principle, the trade balance may be different from zero, indicating a trade surplus or deficit. Over time deficits must be balanced by surpluses and initial net foreign assets. A modified no-Ponzi game condition therefore reads

$$-I_{-1}B_{-1} = \sum_{t=0}^{\infty}[I_0 I_1 ... I_{t-1}]^{-1} P_{Tt} T_t \tag{3.6}$$

However, Fender and Rankin (2008) show that the assumptions of zero initial net foreign assets and the exogeneity of output in the tradables sector together imply that the trade balance is zero at any point in time. This result significantly simplifies the analysis as it allows to abstract from any dynamics introduced by the accumulation of net foreign assets. As in the previous chapter, we differentiate a reference zero inflation steady state (ZISS) around which the equilibrium conditions are linearized from a constant inflation steady state (CISS) in which the economy finds itself one period before the disinflation policy is applied. The log-linear definitions and equilibrium conditions of the open economy model can be found in Appendix A.2.1, where lower-case symbols denote log deviations of variables from their reference steady state values.

Notice that in what follows I will distinguish between the reference zero inflation steady state (ZISS) around which the equilibrium conditions are linearized and a constant inflation steady state (CISS) in which the economy finds itself one period before the disinflation policy is applied. As the economy is closed, bond holdings are zero in both states.[1] The linearized equilibrium conditions are presented in Appendix A.1.1.[2] In the following, lower-case symbols denote log deviations of variables from their reference steady state values, i.e. $v_t = log\frac{V_t}{V_R}$.

3.3 The Role of the Degree of Openness of the Economy

In opening up the model economy I left its basic structure unchanged. This implies that the initial constant inflation steady state (CISS) is the same as in the previous section. I can thus directly proceed to derive the post-disinflation path of the open economy. The disinflation policy is defined in the exact same way as before and is applied to reduce the rate of price inflation from its initial rate μ_I to the new target μ_D in period $t = 0$. I begin the analysis by noticing that price inflation in period t can be expressed as

$$p_t - p_{t-1} = (1 - \alpha\sigma)(e_t - e_{t-1}) + \alpha\sigma(w_t - w_{t-1}) \tag{3.7}$$

[1]While this must necessarily be the case in a closed economy setting, I will impose it upon the open economy by assumption in the next section

[2]The interested reader may refer to Fender and Rankin (2008) for the linearization of the wage setting condition.

This implies that a reduction in CPI inflation can be the result of a reduction in wage inflation, an exchange rate appreciation or both.[3] It suggests that the post-disinflation path of wages may be subject to oscillations of a smaller magnitude than in the closed economy model if the exchange rate adjusts to alleviate them from part of the burden of reducing the inflation rate. In order to investigate this issue, I proceed as in the previous chapter and express the wage setting condition as

$$\begin{aligned} x_t = {} & \frac{1}{1+\beta}[\frac{2\gamma}{1+\gamma}(\frac{1}{1-\alpha\sigma}p_t - \frac{1}{2}\frac{\alpha\sigma}{1-\alpha\sigma}(x_t + x_{t-1})) + \frac{1-\gamma}{1+\gamma}x_{t-1}] \\ & + \frac{\beta}{1+\beta}[\frac{2\gamma}{1+\gamma}(\frac{1}{1-\alpha\sigma}p_{t+1} - \frac{1}{2}\frac{\alpha\sigma}{1-\alpha\sigma}(x_{t+1} + x_t)) + \frac{1-\gamma}{1+\gamma}x_{t+1}] \end{aligned} \quad (3.8)$$

Notice that setting $\alpha = 1$, we are back at equation (2.32), the post-disinflation law of motion in the closed economy setting. The subsequent steps are thus equivalent to the analysis in the previous section. I again define $\phi_t = x_t - p_t$. The law of motion of the open economy in the post-disinflation state is then given by

$$\phi_{t+1} - \frac{1+\beta}{\beta}\frac{1-\alpha\sigma+\gamma}{1-\alpha\sigma-\gamma}\phi_t + \frac{1}{\beta}\phi_{t-1} = \mu_D(\frac{1-\beta}{\beta} - \frac{2\gamma}{1-\alpha\sigma-\gamma}) \quad (3.9)$$

The stability proof in Appendix A.1.2 shows that (3.9) is saddlepoint stable as one of its eigenvalues lies within and the other outside the unit circle.

[3]The reason is that the foreign price of tradables is assumed to be constant.

This result holds for any choice of parameter values within the defined limits. I again use the eigenvalue-eigenvector solution technique of Blanchard and Kahn (1980) to solve the model. As (3.9) is a scalar second order difference equation and its RHS is constant over time, its rational expectations solution is given by

$$\phi_t = \lambda_1 \phi_{t-1} - \mu_D \frac{(1-\beta)/\beta - 2\gamma/(1-\alpha\sigma-\gamma)}{\lambda_2 - 1} \tag{3.10}$$

where λ_1 denotes the eigenvalue that is smaller in absolute value and λ_2 denotes the one that is bigger. Appendix A.1.2 shows that the condition for the presence of oscillations in the post-disinflation state of the economy is now given by $1 - \alpha\sigma - \gamma < 0$. This implies that an economy that is more closed is more likely to be subject to oscillatory behavior after the disinflation policy is applied. The reason is simply that the exchange rate takes a more significant share in the burden of reducing the rate of CPI inflation, the larger is the share of tradable goods in the consumption bundle and the price index.

Figures 3.1 and 3.2 illustrate these results. I have chosen the same parameterizations as in Figures 2.2 and 2.3 except that α now takes the value 0.9 in Figure 3.1 and the value 0.1 in Figure 3.2. A first look at the graphs shows that the exchange rate behaves in a qualitatively equivalent fashion as the new wage. In Figure 3.1 it initially appreciates and then follows an oscillatory path until it converges to its ZISS level. In Figure 3.2, the convergence process is monotonic. The exchange rate appreciates on impact and then gradually depreciates towards its ZISS value. Finally, notice that the oscillations in Figure 3.1 are of a slightly smaller magnitude than in Figure

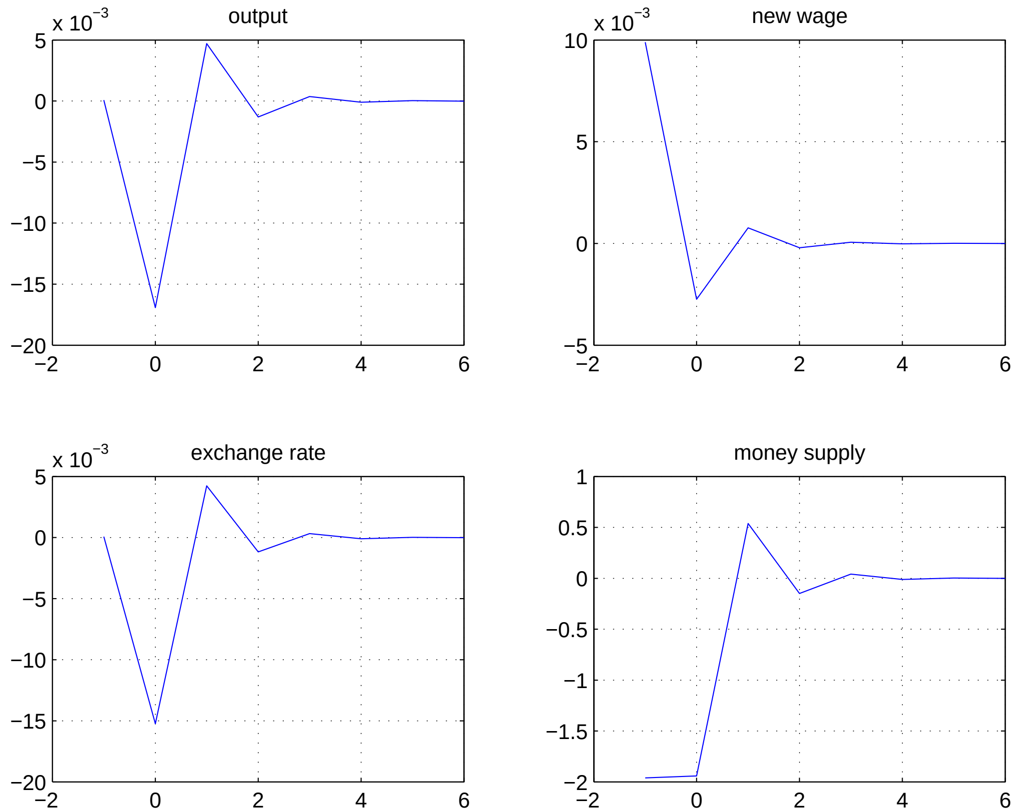


Figure 3.1: Open Economy: Post-Disinflation Path with Oscillations

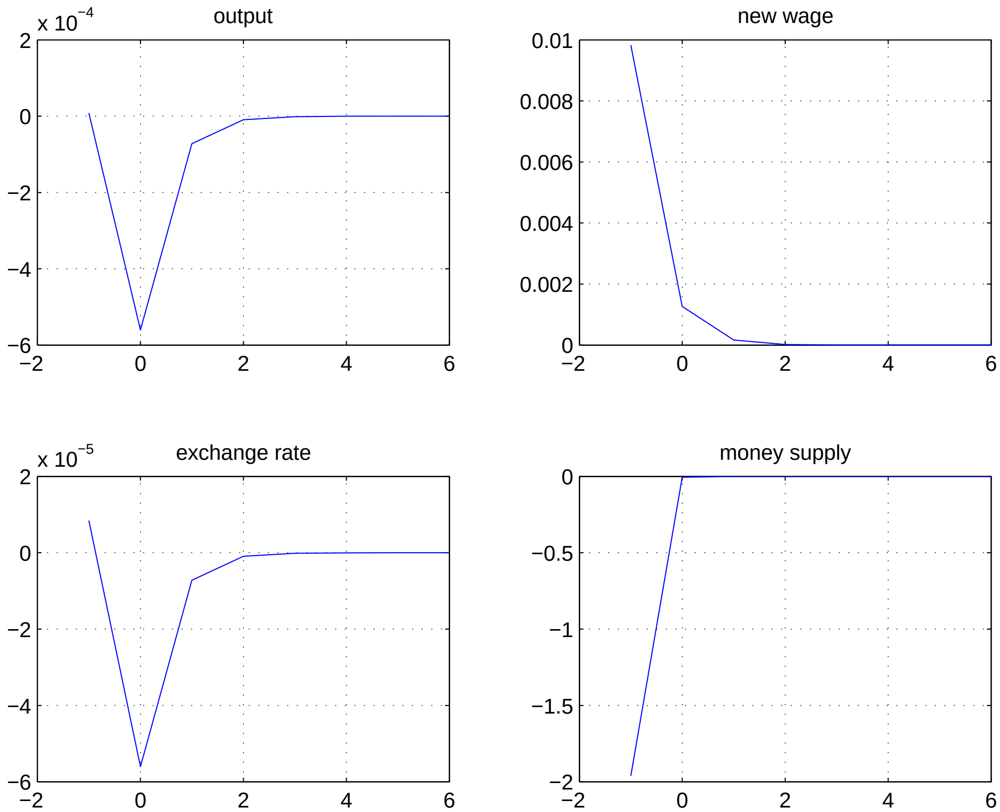


Figure 3.2: Open Economy: Post-Disinflation Path without Oscillations

2.2. This is in line with the intuition that the exchange rate contributes to a smoother post-disinflation path of the economy.

In sum, the exchange rate alleviates wages from part of the burden of reducing the inflation rate when the economy is open. The more open is the economy, the larger is the role played by the exchange rate and the smaller is the magnitude of the oscillations in the post-disinflation state of the economy. Even in the case of constant returns to scale in the production function, the economy always converges to a new steady state and oscillations are never permanent.

3.4 The Nonlinear Open Economy

In this subsection, I investigate the impact of the disinflation policy on the open economy when the model is not linearized. I do so in order to illustrate that the log-linear model hides interesting findings by eliminating the model's non-linearities. In particular, I show that a disinflation from a constant inflation rate to zero cannot be implemented for any realistic parametrization of the model when the returns to scale in the production function are constant and the economy is closed. The reason is that the actions of the policymaker are in this case constrained by a liquidity trap. The case of a partial reduction in inflation is then even more problematic. If inflation is to be reduced to a positive new target, there is no perfect foresight solution to the model at all. However, subsequently I show that these constraints are exclusive to the special case of of a closed economy with constant returns to scale in the production function of its firms.

3.4.1 The Initial Constant Inflation Steady State

As in the previous section, I assume that the economy is initially in a CISS in which real variables are constant and all nominal variables grow at the constant inflation rate $\mu_I = \frac{M_t}{M_{t-1}}$.[4] I again first determine the CISS solution of the model. In order to derive an expression for nontradables output in the CISS, I proceed as before and find that

[4]Note that the definition of μ_I has changed. The definition of μ_D changes accordingly.

$$Y_N = A[(\frac{(\frac{1}{2}+\frac{1}{2}\mu_I^{\varepsilon-1})^{-\zeta}+\beta\mu_I^{\zeta}(\frac{1}{2}+\frac{1}{2}\mu_I^{1-\varepsilon})^{-\zeta}}{(\frac{1}{2}+\frac{1}{2}\mu_I^{\varepsilon-1})^{-1}+\beta(\frac{1}{2}+\frac{1}{2}\mu_I^{1-\varepsilon})^{-1}})^{\frac{1}{\zeta}}(\frac{1}{2}+\frac{1}{2}\mu_I^{\varepsilon-1})^{\frac{1}{1-\varepsilon}}]^{-\sigma} \quad (3.11)$$

where $A = (\frac{1}{\sigma\alpha})^{-\sigma}(\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\sigma\alpha)^{\zeta-1})^{-\frac{\sigma}{\zeta}} > 0$. The rate of money growth enters the expression for nontradables output in a more complex fashion than in the linearized version of the model. I now identify three channels through which the rate of money growth μ_I affects nontradables output in the CISS. These channels will turn out to be important for the adverse effect of the disinflation policy on average output.[5]

(1) *Discounting Channel*: Under wage staggering, the wage that households set in period t has to lie between the ideal wage for period t and the projected ideal wage for period t+1. Under positive inflation, the ideal wage for period t must be lower than the one for period t+1. And as individuals discount future utility according to the discounting parameter $\beta < 1$, the wage set in period t will be set closer to the ideal current wage than to the projected ideal wage for the subsequent period. This, in turn, allows firms to employ more labor at the same cost and increases output. The magnitude of the effect increases the greater is the dispersion between the two ideal wages, i.e. the greater is μ_I. In sum, μ_I has a positive effect on equilibrium output due to its depressing effect on real wages through the discounting parameter β.[6]

(2) *Productivity Channel*: In the presence of inflation, the new wage in

[5]Note that in the linearized version of the model, there is only one, namely the 'discounting channel'.

[6]The smaller is β, the lower will households set their wages.

period t is smaller than the optimal wage in period t+1. Wage staggering then implies that in an arbitrary period t, wages differ between the sector that has just adjusted its wage and the sector whose wage is prevailing from the previous period. This implies that inflation creates wage dispersion between households with the two sectors alternating in setting the higher wage.[7] To be able to infer what wage dispersion implies for equilibrium output, notice that labor skills are imperfect substitutes in the production function. This implies that, ceteris paribus, using equal amounts of each labor types yields a higher average productivity of labor than using unequal amounts. In the presence of wage dispersion, profit maximizing firms substitute labor from the low wage sector for labor from the high wage sector. This choice is optimal under wage dispersion but is inefficient in comparison with conditions under which households in different sectors set the same wages and supply the same amounts of labor and thus attain a higher level of average productivity per unit of labor.[8] In sum, inflation creates wage dispersion which leads to labor substitution and a reduction in productivity. The higher is ε, the less severe is this effect. However, the less severe the effect is, the more do firms engage into labor substitution and the lower is the level of output attained in equilibrium.

(3) *Disutility Channel*: I have elaborated on the fact that wage dispersion induces firms to substitute labor from the low wage sector for labor from the high wage sector. As long as $\zeta > 1$, the resulting intertemporal fluctuation

[7] The exact same degree of wage dispersion that is present in the CISS will be preserved in the post-disinflation state due to the particular nature of the disinflation policy.

[8] These conditions are for instance satisfied in the ZISS around which I have linearized the equilibrium conditions in the previous section.

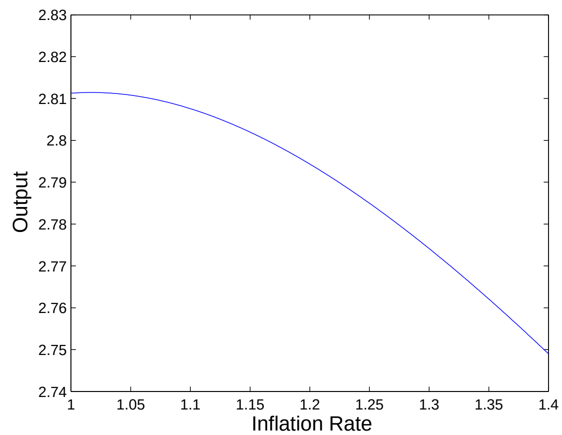


Figure 3.3: The Effect of the Inflation Rate on Output

in the demand for a given household's labor supply increases the disutility the household derives from providing it. This increase in disutility induces households to demand compensation payments in the form of wage increases. The higher the degree of wage dispersion, the greater the increase in the wage level and hence the greater the reduction in output. In sum, μ_I has a negative effect on equilibrium output in the CISS as it increases the overall disutility of work effort. This effect is stronger the greater is ζ.

Figure 3.3 illustrates the effect of μ_I on CISS output. The impact of a marginal increase in inflation varies across different levels of the inflation rate. An increase in the inflation rate affects output positively if the inflation rate is rather low. The reason is that, at low rates of inflation, the 'discounting channel' dominates. The greater is μ_I, the stronger is the degree of wage dispersion and the more do the 'productivity channel' and the 'disutility channel' gain in relative importance. For sufficiently large values of μ_I, the net effect of a marginal increase in μ_I on output is negative. This implies that the distortion resulting from an additional percentage point of inflation worsens with the size of the rate of inflation. The smaller the magnitude of β, ζ and ε, the stronger is the discounting channel relative to the productivity channel and the disutility channel.[9]

I move on to derive the CISS solution for the wage index one period before disinflation. Applying the same procedure as in the previous section, I have that

$$W_{-1} = M_{-1}V[\frac{1}{2} + \frac{1}{2}\mu_I^{\varepsilon-1}]^{\frac{1}{1-\varepsilon}} \tag{3.12}$$

where V is given by

$$V = \frac{1}{Z}[\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\sigma\alpha)^{\zeta-1}\frac{(\frac{1}{2}+\frac{1}{2}\mu_I^{\varepsilon-1})^{-\zeta}+\beta\mu_I^{\zeta}(\frac{1}{2}+\frac{1}{2}\mu_I^{1-\varepsilon})^{-\zeta}}{(\frac{1}{2}+\frac{1}{2}\mu_I^{\varepsilon-1})^{-1}+\beta(\frac{1}{2}+\frac{1}{2}\mu_I^{1-\varepsilon})^{-1}}]^{\frac{1}{\zeta}} \tag{3.13}$$

where $Z_t = \frac{M_t}{S_t}$ and where, assuming that there is no foreign inflation, the money demand function and the interest rate parity condition imply that

$$Z = \frac{1-\delta}{\delta}\frac{\mu_I}{\mu_I - \beta} \tag{3.14}$$

Equations (3.11) and (3.12) determine the state of the economy one period before disinflation.[10]

[9] Ascari (1998) and Graham and Snower (2004) derive the effect of money growth on output in similar frameworks and reach the same qualitative conclusion.

[10] The disinflation policy is unexpected such that the expectation of $\frac{S_0}{S_{-1}}$ as of period $t = -1$ must be μ_I.

3.4.2 The Special Case: Closed Economy with Constant Returns to Scale

I initially focus on the special case of a closed economy and constant returns to scale in the production function. The supply equation (2.3) shows that, as in the linearized version of the model, these assumptions imply that a policy of CPI inflation targeting is equivalent to one of wage inflation targeting. I concentrate on the particular case of $\mu_D = 1$ to begin with. In solving for the post-disinflation path of the economy, I follow the same procedure as in the case of the linearized model.

It is straightforward to determine what the disinflation policy implies for wages set from period $t = 0$ onwards. I use the wage index and impose that the rate of inflation is equal to μ_I until period $t = -1$ and to $\mu_D = 1$ thereafter. I find that in all odd periods starting with period one,

$$X_t = X_{-1} \tag{3.15}$$

and in all even periods starting with period zero,

$$X_t = X_{-2} = \frac{X_{-1}}{\mu_I} \tag{3.16}$$

It is straightforward to show that the wage setting condition (2.10) can, in a complete markets equilibrium, be expressed as

$$X_t = [\frac{\varepsilon}{\varepsilon - 1} \frac{\eta\zeta}{\delta} \frac{W_t^{\varepsilon\zeta} N_t^{\zeta} + \beta W_{t+1}^{\varepsilon\zeta} N_{t+1}^{\zeta}}{W_t^{\varepsilon} N_t / S_t + \beta W_{t+1}^{\varepsilon} N_{t+1} / S_{t+1}}]^{\frac{1}{1+\varepsilon(\zeta-1)}} \tag{3.17}$$

I substitute for the wage index and eliminate employment. This yields

$$X_t = [\frac{\varepsilon}{\varepsilon - 1}\frac{\eta\zeta}{\delta}(\frac{1}{2}X_t^{1-\varepsilon} + \frac{1}{2}X_{t-1}^{1-\varepsilon})^{1-\zeta}\frac{S_t^{\zeta} + \beta S_{t+1}^{\zeta}}{1+\beta}]^{\frac{1}{1+\varepsilon(\zeta-1)}} \tag{3.18}$$

This relationship holds for all $t = 0, 1, 2, \ldots, \infty$. Evaluating (3.18) at each particular time period and using (3.15) and (3.16), I find that

$$S_{t+2} = [[S_t^{\gamma} - (M_{-1}V)^{\zeta}\frac{1+\beta}{\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\frac{1}{2} + \frac{1}{2}B^{1-\varepsilon})^{1-\zeta}}(1 - \beta B^{1+\varepsilon(\zeta-1)})]/\beta^2]^{\frac{1}{\zeta}} \tag{3.19}$$

$$S_{t+2} = [[S_t^{\gamma} - (M_{-1}V)^{\zeta}\frac{1+\beta}{\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\frac{1}{2} + \frac{1}{2}B^{1-\varepsilon})^{1-\zeta}}(B^{1+\varepsilon(\zeta-1)} - \beta)]/\beta^2]^{\frac{1}{\zeta}} \tag{3.20}$$

in odd and even periods respectively, where $B = (\frac{1}{2} + \frac{1}{2}\mu_I^{\varepsilon-1})^{\frac{1}{1-\varepsilon}}/(\frac{1}{2} + \frac{1}{2}\mu_I^{1-\varepsilon})^{\frac{1}{1-\varepsilon}}$.

Equations (3.19) and (3.20) represent the fundamental laws of motion of the economy in the post-disinflation state. In order to learn about the stability properties of the two equations, I differentiate both and evaluate them at their respective steady states. The resulting expressions turn out not to be analytically tractable. I am thus left with the option to establish stability results for given parameter values. I compute the slope of each of the two equations at their respective steady states for various combinations of parameter values within the defined limits. I find a robust result across all parameterizations, namely that both equations exhibit a slope greater than one at their respective steady states and are thus locally unstable. The unique non divergent solution of (3.19) and (3.20) must then be given by their respective steady states. I proceed to rule out unstable solutions as in the previous section. In the post-disinflation state, nominal consumption is thus given by

$$S_t = [(M_{-1}V)^\zeta \frac{1+\beta}{\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\frac{1}{2}+\frac{1}{2}B^{1-\varepsilon})^{1-\zeta}}(1-\beta B^{1+\varepsilon(\zeta-1)})/(1-\beta^2)]^{1/\zeta} \quad (3.21)$$

$$S_t = [(M_{-1}V)^\zeta \frac{1+\beta}{\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\frac{1}{2}+\frac{1}{2}B^{1-\varepsilon})^{1-\zeta}}(B^{1+\varepsilon(\zeta-1)}-\beta)/(1-\beta^2)]^{1/\zeta} \quad (3.22)$$

in odd and even periods respectively. The results show that, as in the linearized version of the model, nominal consumption fluctuates between even and odd periods throughout future periods in the special case of $\mu_D = 1$. The same is true for money supply which is given as a function of nominal consumption by

$$M_t = \frac{1-\delta}{\delta}\frac{S_t}{1-\beta S_t/S_{t+1}} \quad (3.23)$$

However, taking a closer look at (3.22), we observe that a steady state solution for nominal consumption does not exist in even periods for sufficiently large values of μ_I. We know that the initial rate of money growth μ_I is greater than one. Thus, $B < 1$. In fact, B decreases with μ_I. For sufficiently large values of μ_I, the term $B^{1+\varepsilon(\zeta-1)} - \beta$ is negative. This implies that the steady state solution for nominal consumption in even periods does not exist. The reason is the following: nominal consumption is high in odd periods in which wage setters set high wages. It is low in even periods in which wage setters set low wages. A large μ_I implies a high degree of wage dispersion and thus strong oscillations in macro variables. For sufficiently large values of μ_I, the wage set in even periods becomes arbitrarily small and so does nominal consumption. Eventually, there is no positive level of nominal consumption in even periods that is consistent with the wage setting behavior of households and the disinflation policy becomes infeasible. The maximum size of the initial inflation rate μ_I for which the disinflation policy is still feasible in terms of this constraint depends on the choice of parameter values. Unsurprisingly, it hinges crucially on the parameters β, ε and ζ. The reason is that these parameters determine the wage setting behavior of households and thereby the amplitude of the oscillations in macro variables. The greater is the magnitude of each of these parameters, the stronger incentives must be set to induce the required wage setting behavior on the part of the households.

The constraint $B^{1+\varepsilon(\zeta-1)} - \beta \geq 0$ does not, however, turn out to be binding in this model. The reason is that there is a second constraint for the policy to be feasible which takes the form of a liquidity trap. The set of combinations of parameter values that do not violate the second constraint is a strict subset of those that do not violate the first. This implies that there are cases in which the economy hits the liquidity trap although a solution for even periods' nominal consumption exists. In order to fully understand why the economy hits a liquidity trap in this framework, notice that the consumption Euler equation (2.8) is given by

$$I_t = \frac{1}{\beta}\frac{S_{t+1}}{S_t} \tag{3.24}$$

The optimal intertemporal choice of consumption requires the nominal interest rate to be low in odd periods in order to induce a fall in nominal consumption in the subsequent even period. Similarly, a rise in nominal consumption in odd periods requires a high nominal interest rate in the previous even period. The money demand function in equation (2.9) shows that the interest rate increase in even periods causes a fall in even periods' money demand, while the decrease in odd periods causes a rise in odd periods' money demand. The Euler equation and the money demand function taken together imply the following: the greater are the oscillations in nominal consumption that the disinflation policy produces in the post disinflation state, i.e. the smaller is $\frac{S_{t+1}}{S_t}$ in odd periods, the smaller is the nominal interest rate in even periods. As $\frac{S_{t+1}}{S_t}$ approaches β, the nominal interest rate approaches its lower bound $I_t = 1$ and money demand approaches infinity. For $\frac{S_{t+1}}{S_t} < \beta$, the nominal interest rate becomes (notionally) smaller than one and the demand for money flicks from infinity to negative infinity. Mathematically, the reason is that the demand for money is a hyperbola as equation (2.9) shows. Intuitively, the key insight is that the strict disinflation policy hits the lower bound of the nominal interest rate. For a sufficiently strong reduction in inflation, there is no odd periods' nominal interest rate $I_t > 1$ that could induce an upward jump in money demand of the magnitude that is required for the policy to work. The monetary authority is thus constrained by a

Table 3.1: Maximum Value of μ_I Allowing the Policymaker to Avoid the Liquidity Trap

β	ε	ζ	μ_I
0.90	1.4	1.2	1.0052
0.95	3.5	1.4	1.0007
0.95	4.5	1.9	1.0004
0.97	6.0	2.5	1.0001

liquidity trap. In Table 3.1, I list the maximum feasible initial inflation rates that do not violate this feasibility constraint for given parameterizations. It is immediately obvious that the constraint is too strict for the policy to be of any practical relevance in this case.

The above results imply that for sufficiently high initial inflation rates it may not be possible for the monetary authority to set the future path of its policy instrument such that the inflation rate is immediately reduced to zero and the zero inflation rate is sustained throughout future periods. The magnitude of the feasible reductions in the inflation rate is surprisingly small. If one considers the last two cases in Table 3.1 to be realistic calibrations, the analysis implies that the maximum feasible rate of wage inflation which can be immediately and permanently reduced to zero amounts to less than 0.05 percent. The disinflation policy admittedly is very strict. But the results suggest that a complete inflation reduction is infeasible for any practical purposes. An analysis of the effects of the use of strict inflation targeting to reduce the rate of inflation from a positive to a lower positive value therefore suggests itself.

In the framework of the linearized model, I found that the nature of the disinflation policy's impact on the economy in the case of a partial reduction of inflation is equivalent to the particular case of eliminating inflation alltogether. In the nonlinear model, however, it turns out that there is no equilibrium of the model that could support such a policy. In particular, the inflation rate cannot be kept at the new target after disinflation throughout all future periods. Due to the nonlinearity in the wage index formula, the new wage is subject to fluctuations between even and odd periods which grow over time. The new wage set in one of the two sectors grows, while the other falls and eventually hits its zero lower bound. Hence, an equilibrium does not exist. Surprisingly, this is not the case if the inflation rate μ_I is reduced to a new but negative inflation target $0 < \mu_D < 1$. Intuitive explanations as well as mathematical proofs for these findings are presented in Appendix A.1.

I now return to the particular case where $\mu_D = 1$ and proceed to derive the full post-disinflation solution of the model for the case of a reduction of the inflation rate from a positive value to zero. Using nontradables supply and demand, I can express output in the post-disinflation state as

$$Y_{Nt} = [\frac{S_t}{W_{-1}}]^{\sigma} \tag{3.25}$$

Hence,

$$Y_N = [(\frac{1}{2} + \frac{1}{2}\mu_I^{\varepsilon-1})^{\frac{1}{1-\varepsilon}}]^{-\sigma}[\frac{1+\beta}{1-\beta^2}\frac{1-\beta B^{1+\varepsilon(\zeta-1)}}{\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\sigma\alpha)^{\zeta-1}(\frac{1}{2}+\frac{1}{2}B^{1-\varepsilon})^{1-\zeta}}]^{\frac{\sigma}{\zeta}} \tag{3.26}$$

Table 3.2: Nontradables Output for Different Calibrations

Parameters	Cal A	Cal B	Cal C	Cal D
β	0.9	0.95	0.95	0.97
ε	1.4	3.5	4.5	6
ζ	1.2	1.4	1.9	2.5
μ_I*	1.0052	1.0007	1.0004	1.0001

Nontradables Output	Cal A	Cal B	Cal C	Cal D
CISS	0.806029	1.35665	1.12799	1.01772
ZISS	0.805927	1.35664	1.12798	1.01772
Post Disinflation odd	0.843760	1.38503	1.14881	1.02962
Post Disinflation even	0.767502	1.32791	1.10671	1.00556
Post Disinflation avg	0.805631	1.35647	1.12776	1.01759

* $\mu_I = 1$ in ZISS
For all Calibrations: $\eta = 0.1$ and $\delta = 0.7$

$$Y_N = [(\frac{1}{2} + \frac{1}{2}\mu_I^{\varepsilon-1})^{\frac{1}{1-\varepsilon}}]^{-\sigma}[\frac{1+\beta}{1-\beta^2}\frac{B^{1+\varepsilon(\zeta-1)} - \beta}{\frac{\varepsilon}{\varepsilon-1}\frac{\eta\zeta}{\delta}(\sigma\alpha)^{\zeta-1}(\frac{1}{2}+\frac{1}{2}B^{1-\varepsilon})^{1-\zeta}}]^{\frac{\sigma}{\zeta}} \tag{3.27}$$

in odd and even periods respectively.

Table 3.2 shows that post-disinflation nontradables output oscillates between low and high values in a similar fashion as observed in the previous section for the linearized version of the model. Notice that these oscillations are of a surprisingly large magnitude. In case A, in which the rate of wage inflation is reduced by only about 0.5 percent, the magnitude of the oscillations in output is approximately 5 percent. This is empirically absolutely implausible. Average output is moreover smaller in the post-disinflation state compared to both the initial CISS as well as the reference ZISS. The reason is that the positive discounting effect of inflation on output is not present in the POS while the disinflation policy preserves the exact degree of wage dispersion present in the CISS and thus the adverse effects resulting from it.

The presence of wage dispersion in the post-disinflation state also implies that ZISS output is, in contrast to the results in the previous section, higher than average output in the post-disinflation state of the model economy. Finally, notice that ZISS output is almost equal to CISS output for each of the calibrations presented in Table 2. One reason is that the disinflation is of a very small magnitude. Another is that there is no discounting effect but also no wage dispersion in the ZISS. It appears that the positive discounting effect of inflation on output and the adverse effects of wage dispersion almost exactly offset each other in the CISS for the parameter choices presented in Table 2.

In sum, the disinflation policy is infeasible for any practical purposes in a closed economy with constant returns to labor in firms' production functions. For feasible calibrations, it results in a reduction of average output and substantially increases output volatility. The reduction in average output grows, in relative terms, with the magnitude of the initial inflation rate μ_I.[11] In Appendix A.2.3 I show that the disinflation policy, mainly due to the resulting reduction in average output, also reduces welfare if one excludes

[11] Notice that Output in the CISS, ceteris paribus, falls with β, ε and ζ. However, it may increase or decrease with μ_I depending on the reference value of μ_I. Moreover, there is also an interaction effect as the adverse effects of ε and ζ as well as the positive effect of β on output decrease in magnitude the smaller is μ_I. This explains, why it is possible for CISS output to increase from case 1 to case 2 but to decrease from case 2 to case 3 and from case 3 to case 4.

money balances from the utility function.

3.4.3 The General Case: Open Economy and Decreasing Returns to Scale

The purpose of the present section is to show that both a complete and a partial disinflation of an empirically relevant size are possible in the nonlinear model if one abstracts from the special case of a closed economy and constant returns to scale. The post-disinflation path of the economy moreover looks strikingly similar to the equivalent case in the log-linear model.

As before I solve for the post-disinflation path of the economy by focusing on its law of motion, i.e. the wage setting condition, which can be expressed as

$$X_t = [\frac{\varepsilon}{\varepsilon - 1}\frac{\eta\zeta}{\delta}\frac{W_t^{\varepsilon\zeta}Y_{Nt}^{\zeta/\sigma} + \beta W_{t+1}^{\varepsilon\zeta}Y_{Nt+1}^{\zeta/\sigma}}{W_t^{\varepsilon}Y_{Nt}^{1/\sigma}/S_t + \beta W_{t+1}^{\varepsilon}Y_{Nt+1}^{1/\sigma}/S_{t+1}}]^{\frac{1}{1+\varepsilon(\zeta-1)}} \tag{3.28}$$

As in the case of the linearized model, we need to normalize all nominal variables by the price level before we can solve the model. For each nominal variable Q_t, I define $\tilde{Q}_t = \frac{Q_t}{P_t}$. Using this definition as well as the fact that $P_{t+1} = P_t\mu_D$ in the post-disinflation state, it is straightforward to show that the following system of equations determines the post-disinflation path of the economy:

$$\tilde{X}_t = [\frac{\varepsilon}{\varepsilon - 1}\frac{\eta\zeta}{\delta}\frac{\tilde{W}_t^{\varepsilon\zeta}Y_{Nt}^{\zeta/\sigma} + \beta(\tilde{W}_{t+1}\mu_D)^{\varepsilon\zeta}Y_{Nt+1}^{\zeta/\sigma}}{\tilde{W}_t^{\varepsilon}Y_{Nt}^{1/\sigma}/\tilde{S}_t + \beta(\tilde{W}_{t+1}\mu_D)^{\varepsilon}Y_{Nt+1}^{1/\sigma}/(\tilde{S}_{t+1}\mu_D)}]^{\frac{1}{1+\varepsilon(\zeta-1)}} \tag{3.29}$$

$$\alpha^{\alpha} = \tilde{P}_{Nt}^{\alpha}\tilde{S}_t^{1-\alpha} \tag{3.30}$$

$$\tilde{W}_t = [0.5\tilde{X}_t^{1-\varepsilon} + 0.5(\frac{\tilde{X}_{t-1}}{\mu_D})^{1-\varepsilon}]^{\frac{1}{1-\varepsilon}} \tag{3.31}$$

$$Y_{Nt} = (\frac{\sigma\alpha\tilde{S}_t}{\tilde{W}_t})^{\sigma} \tag{3.32}$$

$$\tilde{P}_{Nt} = (\alpha\tilde{S}_t)^{1-\sigma}(\tilde{W}_t/\sigma)^{\sigma} \tag{3.33}$$

We need to resort to numerical methods to solve this system of equations. A set of initial conditions can be derived based on equations (3.11) and (3.12). The latter define the state of the economy one period before disinflation. As a solution procedure, I use the Newton-type algorithm first proposed by Laffargue (1990).[12]

Figures 3.4 and 3.5 consider a disinflation from two percent down to price stability. The solid lines represent the post-disinflation path of the economy in the log-linear version of the model using the same parameterizations as in Figures 2.2 and 2.3.[13] The dashed lines represent the corresponding path of the economy in the full nonlinear model. It is immediately obvious that a disinflation of a realistic size is indeed possible in the nonlinear version of the model. Moreover, not only is the post-disinflation path of the model economy qualitatively equivalent to the case of the log-linear version of the model, it is

[12] The algorithm is employed in Dynare for the solution of deterministic models.

[13] I converted the values from log deviations back to level terms.

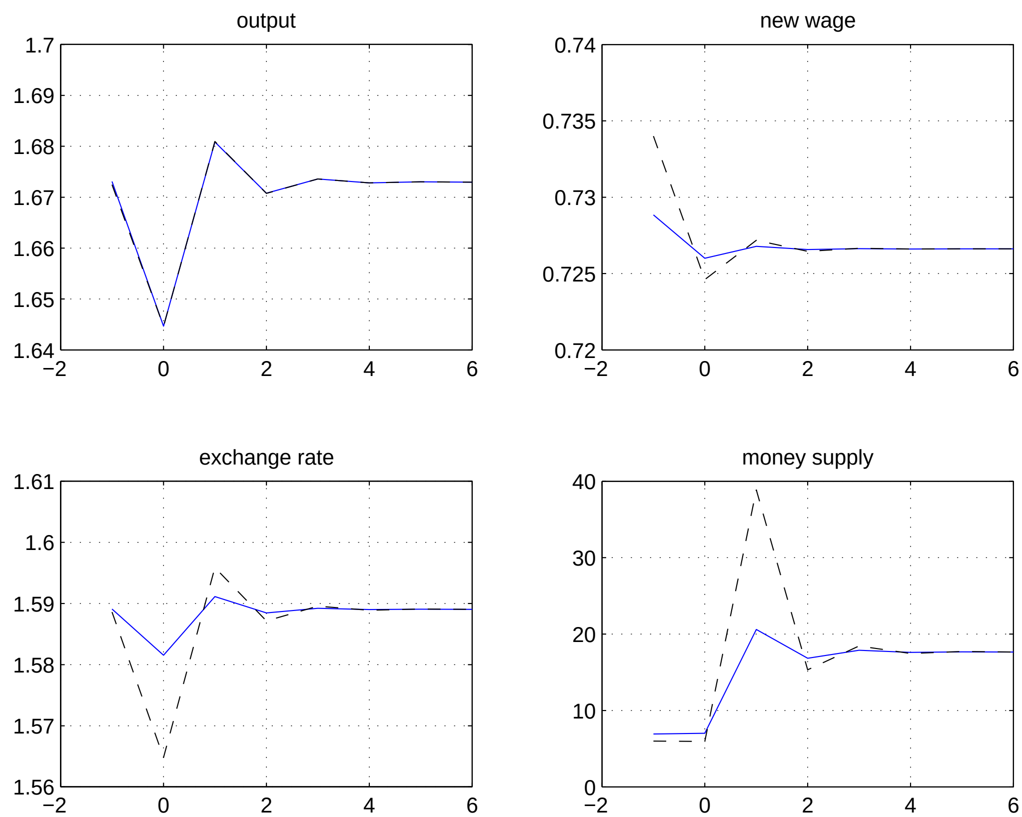


Figure 3.4: Comparison of the Linear and the Nonlinear Model I

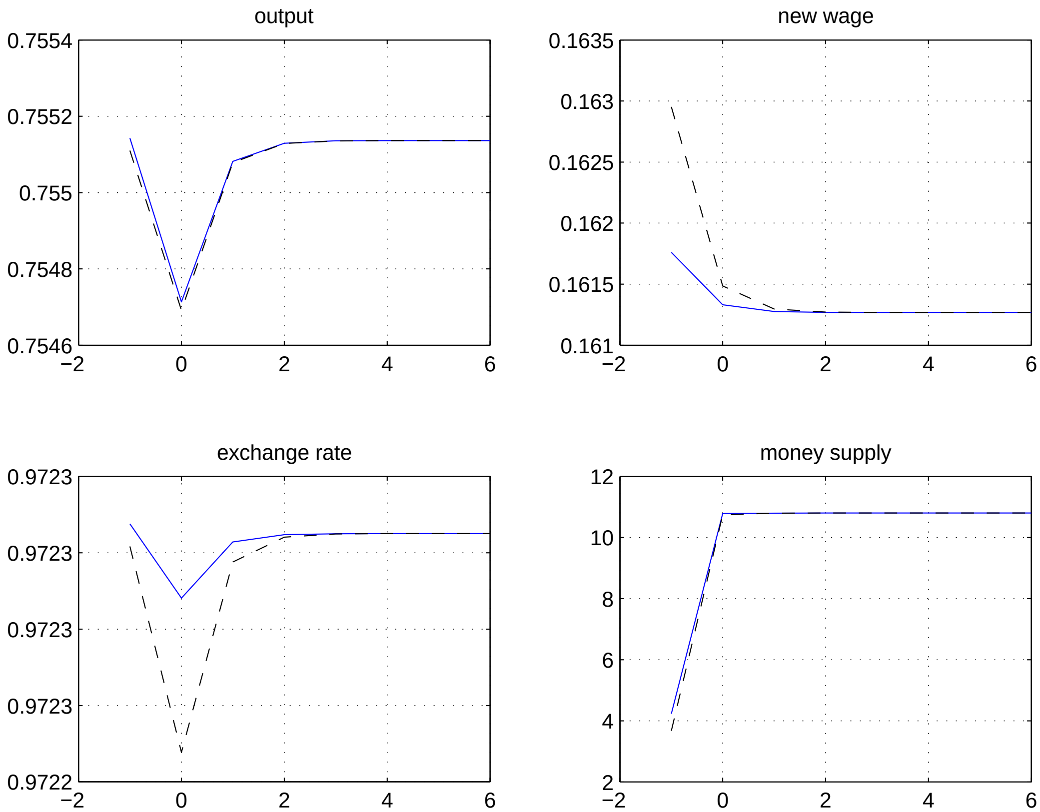


Figure 3.5: Comparison of the Linear and the Nonlinear Model II

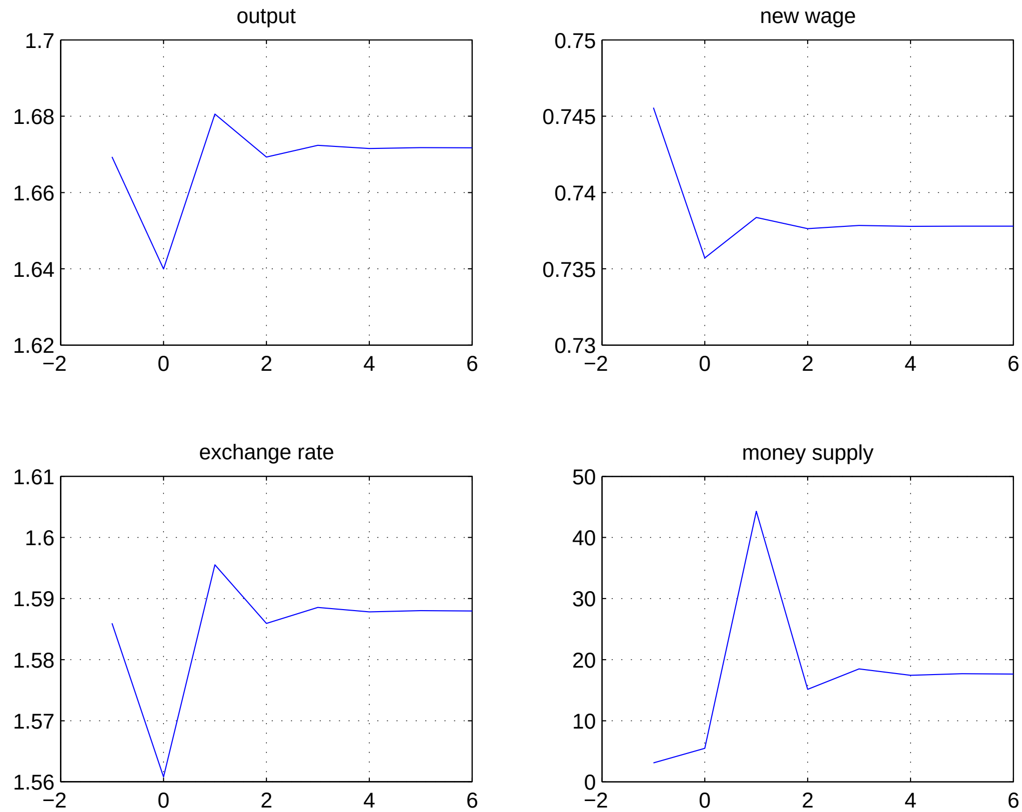


Figure 3.6: Nonlinear Model: Partial Disinflation with Oscillations

even quantitatively very similar. While the magnitude of the fluctuations in nominal variables is greater in the nonlinear model, the movements in output are almost precisely of the same magnitude. This implies that we can draw the same conclusions from these results as I did in the previous section for the log-linear version of the model.

Figures 3.6 and 3.7 present the results from reducing the rate of price inflation from 5 percent to 3 percent in the nonlinear model. The nominal variables are shown as normalized by the price index. It is immediately obvious that, contrary to the special case of a closed economy and constant returns to scale, not only disinflations of an empirically relevant size but

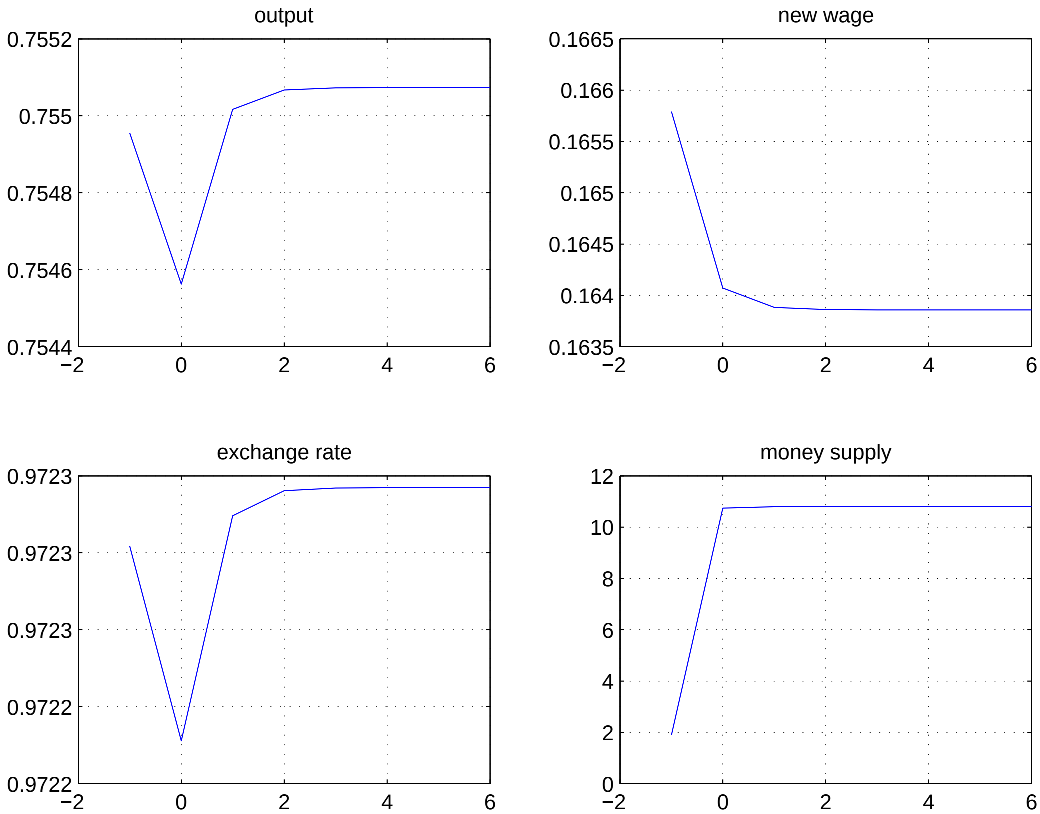


Figure 3.7: Nonlinear Model: Partial Disinflation without Oscillations

also partial disinflations are possible in the nonlinear model. As in the log-linear version of the model, the qualitative conclusions arising from partial disinflations are the same as in the case of a complete disinflation to price stability.

3.5 Discussion

This chapter investigated how our previous conclusions about the impact of the disinflation policy change when the model economy is extended to an open economy setting. In addition to the returns to scale in the production function, the degree of openness of the economy was identified as a decisive factor in determining the presence and the magnitude of the oscillations in the post-disinflation state of the economy. I showed that the degree of openness matters because the exchange rate acts as a stabilizer along the post-disinflation path of the economy by effectively relieving wages of part of the burden of reducing the inflation rate. The more open is the economy, the greater is the share of the burden it can successfully manage.

Furthermore, I investigated the impact of the disinflation policy on the economy in the framework of the full nonlinear model. Although the results are in general qualitatively very similar to the case of the loglinear version of the model, a set of additional conclusions emerges in the special case of a closed economy with constant returns to scale in firms' production functions. It turns out that the policymaker faces a surprisingly strict feasibility constraint that does not allow for the policy to be carried out in virtually any case of empirical relevance.

Chapter 4

Bank of England Interest Rate Announcements and the Foreign Exchange Market

4.1 Introduction

The Bank of England (BoE) was granted operational independence to set its key policy interest rate by the incoming UK Labour government in May 1997, with the goal of creating policy consistent with stable inflation and economic growth.[12] In practice, interest rate decisions are made by the Bank's Monetary Policy Committee (MPC), which meets for two days each month-as well as an additional pre-meeting briefing day-and issues a statement regarding interest rate decisions at noon on the second meeting day. This framework

[1]This chapter is co-authored with Michael Melvin, Michael Sager and Mark P. Taylor.

[2]Prior to August 2006, policy decisions were framed in terms of the repurchase, or repo, rate. We use the names Bank rate and repo rate interchangeably.

allows a natural laboratory setting for examining the impact of monetary policy decisions around a known time and date. Since market participants know that interest rate announcements arrive at noon on the second meeting day, there may be positioning prior to the announcement and news effects after the announcement that result in systematic patterns in exchange rate behavior on MPC meeting days that differ from other days. A stated aim of the new policy regime was that monetary policy should be more transparent than hitherto (King, 2000). The availability of the record of MPC decisions therefore affords us a rare opportunity to examine how the decisions of the key policy-setting committee are impounded into financial prices. In this paper, we concentrate on an examination of the pattern of exchange rate volatility surrounding the MPC's interest-rate decisions as well as the role played by the surprise content in the announcements.

Since activities directly related to each MPC meeting are spread over three different days (see Section 4.3, below), our analysis will include an examination of the pre-meeting briefing day, the first day of the meeting, and the second day of the meeting when the policy decision is made and publicised (as well as days unrelated to the meetings, to serve as controls). Both daily and high-frequency, intraday data are employed in the analysis. The daily data provide a bird's eye view of market behavior around MPC meetings, using a generalised autoregressive heteroscedastic (GARCH) framework. Given the findings of this low-frequency analysis, a microscope is then taken to the data to examine exchange rate dynamics on days related to MPC meetings. The intraday econometric framework is provided by a Markov switching model where exchange rate returns switch between a high-volatility, informed-trading state, and a low-volatility, uninformed or liquidity trading state. A key difference from the usual Markov switching model employed in financial analysis is our incorporation of endogenous shifts in the transition probabilities, where these shifts are modeled as a function of variables related to the MPC meeting and policy outcomes.

We choose to employ a Markov switching framework in order to allow for an alternative characterization of macroeconomic news effects on the foreign exchange market. The underlying hypothesis is that macroeconomic news do not simply affect the market as shocks to otherwise continuous processes. On the contrary, news effects may change the entire data generating process for a financial variable. One reason is that "hot-potato" trades are likely to dominate the market to an unusual degree as dealers adjust their inventory and offload onto other dealers, effectively generating a multiplier effect on trades (Lyons, 1994). It is difficult to believe that this adjustment period is characterized by the same data generating process that governed the market prior to the news impact. An econometric specification allowing for regime switches therefore appears appropriate. Indeed, one particular benefit of applying such a model is that it facilitates a plausible interpretation of observed nonlinearities. Moreover, and in contrast to the deterministic models typically employed in similar analyses, the framework allows for a probabilistic and thus very flexible characterization of the data. In particular, by modeling switching probabilities endogenously, we allow the probability of regime switching to vary at various points during MPC meeting days, rather than modelling the switch deterministically. Given the notoriously capricious nature of financial markets, our approach therefore provides an interesting alternative perspective on news effects on financial markets.

The next section provides a brief review of the literature on the financial effects of macroeconomic news announcements. In Section 4.3 we provide some background institutional details on the MPC and the UK monetary policy-setting process. Section 4.4 contains a discussion of our econometric methodology and the various hypotheses to be tested. Section 4.5 describes our data sets and contains our main empirical findings. Finally, Section 4.2 summarises our conclusions and discusses directions for future research.

4.2 Exchange Rate and Asset Price Effects of Monetary Policy Announcements: A Brief Review of the Literature

Early intraday studies of macroeconomic news effects on exchange rates, such as Hakkio and Pearce (1985) and Ito and Roley (1987), tend to provide mixed results in terms of the significance of news announcements on exchange rate movements. One possible reason for this finding was the coarseness of the sampling intervals, with observations of exchange rates taken at opening, noon and closing. Clearly, if news effects work themselves out within periods less than several hours, then observing the market at three equally spaced points over the trading day will miss much of the action. The increased availability of high frequency intraday foreign exchange rate data during the 1990s considerably advanced research in this area.

Intraday exchange rate volatility effects of news announcements were first documented by Ederington and Lee (1993, 1995, 1996).[3] Ederington and Lee (1993) use 5-minute tick data from November 1988 to November 1991 for mark-dollar, as well as various interest rate futures, and report conclusions consistent with our findings below. They estimate a series of regressions of the deviation of the absolute value of exchange or interest rate returns in a given five minute period on day j from the average return during that period across the whole sample as a function of a series of dummy variables that designate the publication schedule of various US macroeconomic data series. Ederington and Lee (1993) conclude in favour of a significant change in intraday exchange and interest rate volatility upon publication of various macroeconomic series, including the monthly employment report, producer price inflation and trade data, with the standard deviation of five minute returns immediately after publication at least five times higher on announcement days than on non-announcement, or control, days. Ederington and Lee (1993) also find that although the greatest volatility impact occurs within one minute of publication, the standard deviation of returns remains significantly above normal for up to forty five minutes after publication for a number of macroeconomic series.

In an extension to their original paper, Ederington and Lee (1995) perform a similar analysis using 10-second data, and conclude that much of the price reaction to macroeconomic news is actually completed after only 40 seconds. They also find evidence of a pre-announcement volatility effect immediately ahead of key macroeconomic data releases, consistent with

[3] Taylor (1987, 1989) provides early, high-frequency studies of the foreign exchange market and finds some evidence of the impact of news on deviations from covered interest rate parity.

our findings below. Similarly, Ederington and Lee (1996) report significant volatility effects from macroeconomic data releases in the interest rate options market, although they find against any such effect in mark-dollar option volatility.

A number of papers have since reported findings similar to Ederington and Lee (1993), for both macroeconomic data releases and monetary policy announcements and statements. These include Andersen and Bollerslev (1998), in the context of a wider study of the determinants of mark-dollar volatility, and Goodhart, Hall, Henry and Pesaran (1993). Goodhart et. al. apply a GARCH-M methodology to sterling-dollar tick data over the period April to July 1989 to analyze the volatility impact of an announced BoE interest change and publication of US trade data, both of which occurred in May 1989. Their findings are generally consistent with ours reported below, in that they find significant evidence of a non-permanent volatility impact due to the monetary policy announcement and US trade data publication. They find this volatility effect to be more persistent than either our results or those of Ederington and Lee (1993), and suggest that it remains in the data during the subsequent 4-5 days. Almeida, Goodhart and Payne (1998) perform a similar high frequency analysis of the volatility impact of US and German macroeconomic data releases using five minute tick data for mark-dollar over the sample period January 1992 to December 1994. They, too, find evidence of non-permanent volatility effects. Their conclusion that these effects generally dissipate within fifteen minutes of publication for US data releases, and approximately three hours for German releases, are more consistent with our findings below. Although fewer German data releases examined by Almeida et. al. have a significant impact upon the volatility of exchange rate returns than US series, the number of significant German releases increases when the authors account for the proximity of the next Bundesbank policy meeting; the closer to this meeting, the more likely the Bundesbank will act upon any surprises contained in data releases.

Faust, Rogers, Swanson and Wright (2003) use intraday, daily and monthly data from 1994 to 2001 to estimate structural vector autoregressions (SVARs) incorporating current and future US and foreign short-term interest rates and exchange rate series in order to assess the contemporaneous effect of a US monetary policy shock on other variables in the SVAR.[4] Although the results for future interest rates are mixed, the impact of the monetary shock on both exchange rates using high frequency data is positive (meaning that a surprise rate increase depreciates the value of the dollar) and statistically significant. In a similar vein, Harvey and Huang (2002) examine the impact of Federal Reserve open market operations on a range of interest and exchange rates using GMM estimation and intraday data - specifically, two-minute and hourly returns - over the period 1982 to 1988. In this case, though, while the authors find in favour of a significant increase in intraday interest rate futures volatility associated with so-called Fed Time, they conclude against any significant, generalised increase in exchange rate return volatility.[5]

In a complementary study, Andersen, Bollerslev, Diebold and Vega (2003) focus on detecting shifts in the conditional mean rather than the volatility

[4]Interest rates are measured using futures contracts for Eurodollar, Libor and Fibor/Euribor. Exchange rates included are sterling and mark/euro, both expressed in terms of the US dollar.

[5]They also find evidence that interest rate volatility is actually greater when the Fed does not conduct operations during the allotted time than when it does.

of exchange rates. Using five-minute tick data for the Swiss franc, mark, sterling and yen, all expressed in terms of the US dollar. Andersen et. al. examine the impact of Federal Reserve policy announcements, as well as a variety of macroeconomic data series from the US and Germany, over the sample period January 1992 to December 1998. The authors find in favor of a significant, asymmetric jump effect associated with shocks due to policy announcements by the US Federal Open Market Committee (FOMC) and a number of US data releases immediately following publication of many data series; negative US data surprises often exhibited a larger impact upon exchange rates than positive surprises. By contrast, and yet consistent with the findings of Almeida et. al. (1998), only relatively few German data releases exert a statistically significant effect upon exchange rate levels.

A number of studies complementary to our research have analysed the volatility impact of monetary policy announcements, as well as statements and speeches by central bank officials, using daily data. These include Gurkaynak, Sack and Swanson (2005), Reeves and Sawicki (2005), Kohn and Sack (2003), Ahn and Melvin (2007) and Ehrmann and Fratzscher (2007).

Gurkaynak, Sack and Swanson (2005) analyze the effect of Fed Funds rate changes and the accompanying FOMC policy statements on bond yields and stock prices using factor analysis. They find that two latent factors are necessary in order to capture the asset price effects of monetary policy, with the former associated with the interest rate change and the latter associated with the FOMC statements. The strong policy implication of this research is, therefore, that both monetary policy actions and statements may have important effects on asset prices.[6]

Reeves and Sawicki (2005) analyze the impact on three-month forward interest rates and long gilt futures of the publication of MPC minutes[7], the quarterly BoE Inflation Report, as well as MPC member speeches and regular testimonies to parliamentary committees. Using both daily and intra-day observations over the period June 1997 to December 2005, Reeves and Sawicki conclude in favour of a significant interest rate volatility effect due to the publication of MPC minutes and the Inflation Report (although in this case only using intra-day data). Kohn and Sack (2003) perform a similar analysis on the volatility impact of policy statements by the FOMC, as well as congressional testimony and speeches by former Chairman Greenspan for the trade-weighted dollar, a range of interest rates and S+P500 returns using daily data over the sample period January 1989 to April 2003. They find that FOMC statements generate a significant change in the volatility of interest rates, but no significant change in the volatility profile of either the dollar or S+P500 returns. This finding is consistent with the evidence that we present below using daily data.

Ahn and Melvin (2007) conduct an intradaily examination of exchange rate regime switching for Federal Reserve FOMC meeting days and find surprising evidence of switches to a high-volatility informed trading state during the time of the meeting rather than at meeting end when decisions are announced. An extensive search of public news suggests that this informed

[6]Grkaynak, Sack and Swanson (2005) find, in fact, that policy statements have a much greater impact on longer-term Treasury bond yields.

[7]Since October 1998, MPC minutes are published thirteen days after the associated policy announcement that is the focus of our study.

trading state cannot be explained as the response to public information. This is consistent with a market where informed traders are taking positions in advance of the meeting end based upon their expectations of the outcome.

Finally, Ehrmann and Fratzscher (2007) undertake an EGARCH study of Federal Reserve, BoE and ECB monetary policy announcements and broader statements regarding economic outlook using daily data over sample periods that begin in 1997 for the BoE, and 1999 for both the Federal Reserve and ECB; all sample periods run until 2004. Although evidence for exchange rates is mixed, Ehrmann and Fratzscher do conclude that policy announcements by all three central banks exert a significant impact upon the volatility of interest rates. In addition, the impact of BoE policy announcements is significantly larger than either the Federal Reserve or ECB. This second finding is consistent with the authors' hypothesis that the BoE combination of collegial communication strategy and individualistic voting strategy leads to more regular and significant policy announcement shocks than for either the Fed or ECB. The volatility impact of broader statements on economic outlook is only significant in the case of the Federal Reserve.

4.3 The Monetary Policy Committee

In May 1997 Gordon Brown, then UK Chancellor of the Exchequer, announced that the BoE would be given operational responsibility for setting interest rates via the newly created MPC.[8] The MPC was to focus on an inflation target of 2.5 percent on a two-year horizon for the retail price index

[8] For institutional background on the MPC and the monetary policy process, see Bean (1999). Note that inflation targeting had been adopted in the UK since 1992.

excluding mortgage interest payments.[9] Conditional on maintenance of the inflation target, the MPC could also address fluctuations in economic growth and employment.

The MPC is comprised of nine members. Five are drawn from the BoE: the Governor, the two Deputy Governors, and two Executive Directors. The other four members are drawn from outside the Bank and are appointed by the Chancellor of the Exchequer. At the time this paper was written, the four external members included two academic economists and two business economists. The Governor serves as the Committee chair.

The Committee meets monthly, normally on the Wednesday and Thursday following the first Monday of each month. The meeting dates for each year are published well in advance of the meetings.[10] On the Friday morning prior to each meeting, the Committee meets for a briefing to prepare for the meeting. Summaries of important news and trends are provided by senior BoE staff. On the Monday and Tuesday prior to the meeting, the BoE staff prepares any additional background information and analysis required by the Committee. On these days MPC members receive written answers to any questions that arose at the Friday briefing along with any new data releases or important news.

The monthly MPC meeting typically begins at 3.00 pm on Wednesday afternoon with a review of the state of the UK and world economy. The BoE Chief Economist starts the meeting with a short summary of any major events since the Friday briefing. On Thursday morning, the MPC reconvenes and

[9]This policy goal was subsequently changed to 2.0 percent in December 2003, and is now defined in terms of the harmonized consumer price index.

[10]These are published at www.bankofengland.co.uk.

the Governor begins with a summary of the major issues. Members are then invited to state their views of the appropriate policy to follow. The Deputy Governor responsible for monetary policy will usually speak first with the Governor speaking last. Ultimately, the Governor offers a motion that he suspects will result in a majority vote and then calls for a vote. Members vote with a one-member, one-vote rule. Those in the minority are asked to state their preferred level of interest rates. Lastly, the press statement is developed. If the decision is to change interest rates or follow a policy that was not expected by the market, the press statement will include the reasons for the action taken. In other cases, simply the decision is reported. This decision is announced at noon, London time. Following the announcement, policy is implemented with open-market operations beginning at 12:15 pm.

4.4 Methodology

The focus of this paper is on inference regarding movements in the dollar-sterling exchange rate during MPC meetings. Given that the foreign exchange market knows when the MPC meets and when its decisions are announced, we want to examine evidence regarding any market positioning before and during the meeting and as to whether these effects are driven by the news content of the respective policy announcement.

A logical first step is to examine whether meeting days are different from other days as well as from one another in terms of systematic patterns in dollar-sterling exchange rate movements. As discussed above, given the multi-day structure of MPC deliberations, one may hypothesise that the foreign exchange market forms an opinion about the likely meeting outcome prior to the public announcement at noon on the second day of the meeting. This does not have to rest upon information leaks from the Committee. It may be that traders close down trade positions in advance of the interest rate decisions in order to limit their risk exposure precisely because they are unsure about the upcoming announcement. Furthermore, such behavior may be driven by astute MPC-watchers' informed opinions of the likely Committee vote. An analogy in the Federal Reserve case is the often-cited story of how Fed-watchers at one time gauged the likely FOMC decision by the size of the briefcase that former Chairman Alan Greenspan carried to work. The idea was that a thick briefcase signaled a likely interest rate shift while a thin briefcase signaled a high probability of no change in policy. No doubt, there are many such stories one could gather from MPC watchers as well.

We explore the evidence in the data regarding briefing days, first meeting days, and second meeting days by initially analyzing daily returns for USD/GBP. We estimate simple linear models of daily exchange rate returns incorporating dummy variables for days of MPC briefings, first, and second meeting days as well as a variable indicating the size of the interest rate change:

$$\Delta e_t = a + b_0 Briefing + b_1 Day1_t + b_2 Day2_t + b_3 \Delta i_t + \varepsilon_t \tag{4.1}$$

where Δe_t is the change in the logarithm of the exchange rate on day t, and *Briefing*, *Day1*, and *Day2* are dummy variables equal to 1 on the respective MPC meeting day and equal to zero otherwise. These dummy variables are subsequently incorporated into the conditional variance equation of a GARCH model. Following Andersen and Bollerslev (1997) and Jones, Lamont and Lumsdaine (1998), we allow the announcement effects to have a temporary impact on the conditional variance only, on the basis that announcement effects are likely to die out in less than a full day:[11]

$$\Delta e_t = \mu + c_1 \Delta e_{t-1} + ... + c_n \Delta e_{t-n} + \sqrt{s_t}\varepsilon_t \tag{4.2}$$

$$\varepsilon_t | \Omega_{t-1} \backsim N(0, h_t) \tag{4.3}$$

$$s_t = (1 + \delta_0 Briefing + \delta_1 Day1 + \delta_2 Day2) \tag{4.4}$$

$$h_t = \omega + \alpha \varepsilon_{t-1}^2 + \beta h_{t-1} \tag{4.5}$$

where Ω_{t-1} denotes the information set at time t-1. The conditional variance on any given day t is therefore given by $s_t h_t$ and, e.g., by $(1 + \delta_2)h_t$ on second meeting days. This implies that δ_2 captures the percentage increase in the conditional variance on second meeting days. Estimation of the model is carried out using quasi-maximum likelihood. The results of

[11]We thank an anonymous referee for suggesting this specification.

this daily analysis can help inform us as to whether exchange rate returns and their volatility differ around the time of MPC meetings, or in between the various days of these meetings, and according to whether the respective policy announcements on second MPC meeting days came as a surprise to the market.

We then take a microscope to the data for second meeting days to examine the intraday behavior of returns on days when a policy decision is announced. Before turning to the questions to be examined, the econometric framework employed in our intraday analysis is introduced.

It is usual to think of high-frequency exchange rate data on any given day as bounded within a fairly narrow band and exhibiting first-order autocorrelation. By contrast, on MPC meeting days we may expect important news to be received by the market. We find it convincing to think of these news effects as changing, temporarily, the entire data generating process of the exchange rate - and other financial variables - rather than simply introducing a one-time shock to an otherwise continuous process. Intuitively, so-called "hot-potato" trades are likely to dominate the market to an unusual degree in the immediate aftermath of the news as dealers adjust their inventory and offload onto other dealers, effectively generating a multiplier effect on trades (Lyons, 1994). It is difficult to believe that this adjustment period is characterised by the same data generating process that governed the market prior to the news announcement. An econometric specification allowing for regime switches therefore appears appropriate. We chose to adopt the Markov switching framework associated with Hamilton (1990, 1994), allowing the switching probabilities to be endogenously determined (Diebold, Lee and Weinbach, 1994).

An important advantage of this framework is that it facilitates a plausible interpretation of observed nonlinearities and allows for probabilistic rather than deterministic switching between regimes.[12]

A Markov-switching first-order autoregressive model for exchange rate returns is postulated as follows:

$$\Delta e_t = \mu(S_t) + \rho(S_t)[\Delta e_{t-1} - \mu(S_{t-1})] + \varepsilon_t \tag{4.6}$$

where $\varepsilon_t \backsim N[0, \sigma^2(S_t)]$ and where Δe_t is the change in the logarithm of the exchange rate at time t. Note that the mean of the exchange rate returns process μ, the autocorrelation coefficient ρ, and the variance of the innovation, σ^2 are allowed to take on one of two values depending on the realization of an unobserved state variable $S_t \epsilon [1, 2]$. In our application, we assume a two-state Markov process. One of the states (say, state 2) may be thought of as reflecting the usual pattern of exchange rate returns with negative autocorrelation and a relatively small variance. This tranquil state is the normal state that would be associated with liquidity trading when no important information arrives in the market. The other state (say, state 1) may be thought of as the informed-trading state when volatility is high and

[12]Along with our hypothesis that intraday news does not generate a one-time shock to the distribution of variables, as discussed above, the opportunity to interpret nonlinearities is a principal motivation for employing the Markov switching framework in our intraday analysis rather than simply continuing with the GARCH analysis used at the daily frequency.

realized returns much larger than normal (Easley and O'Hara, 1992; Lyons, 2001).

Thus far, our proposed methodology is similar to that employed, inter alia, in Engel and Hamilton (1990). However, we diverge from the traditional Markov approach by modelling the probability of switching from one regime to another endogenously. Denoting the transition probability of switching from regime j to regime i at time t as P_t^{ij} for $i, j\epsilon[1, 2]$, we can write the postulated functions for the transition probabilities, conditional upon information at time t, I_t, and the previous state, as

$$P_t^{ii} = Pr[S_t = i | S_{t-1} = i, I_t] = \Phi[\alpha_{ii} + \beta_{ii}^{'} X_t] \tag{4.7}$$

where Φ denotes the cumulative normal density function (in order to ensure that the probabilities lie in the unit interval) and where $X_t \epsilon I_t$ is a vector of variables known at time t which may influence the transition probability according to the vector of loadings β_{ii}. Given P_t^{11}, we implicitly have $P_t^{21} = 1 - P_t^{11}$. Similarly, given an estimate of P_t^{22}, we implicitly have $P_t^{12} = 1 - P_t^{22}$.

The Markov-switching framework is applied to high-frequency data to address several questions of interest in the intraday setting. First, can we identify endogenous regime switching? Are the transition probabilities driven by the news component in the policy announcements? To test if the MPC policy announcement released at noon on the second meeting day is price-relevant public news, we incorporate various dummy variables in the explanatory variable vector X_t. These dummies were set equal to one for a certain afternoon period, say noon to 13:00, and to zero otherwise.

A second question of interest is whether evidence exists of positioning during the second meeting day prior to the policy announcement at noon? To address this question, we incorporate dummy variables equal to one for various time intervals prior to noon and zero otherwise. We explore alternative definitions over different morning time intervals as a sensitivity analysis.

4.5 Data and Empirical Findings

Our data sample spans over a decade, running from the inception of the Monetary Policy Committee in June 1997 through to October 2007, and incorporates 126 MPC meetings. Tables A.2 to A.4 in the Appendix list the MPC meeting days in our sample and the associated interest rate decision for each meeting. We classify an MPC decision as a surprise to the market if it differs from the median expectation taken from a survey of market economists by Bloomberg.[13] The standard deviation of analysts' expectations is reported as a measure of forecast dispersion. The tables also provide a range of alternative surprise measures - based on 3 months short term interbank rates (IB) as well as 3 months sterling interest rate futures contracts

[13]This survey is carried out on the Friday before each MPC meeting (i.e. on the same day as the pre-briefing of the MPC by staff members of the BoE) and asks respondents for the magnitude-if any-of the interest rate change that they expect to result from the upcoming meeting. In its current guise, the survey collates the expectations of up to 60 financial economists. Although the sample of economists is not necessarily the same from one month to the next, a core subset ensures continuity and the survey is in any case is designed to capture market expectations.

on the London International Financial Futures Exchange (LIFFE) - to be used for robustness checks.[14]

Tables A.2 to A.4 suggest that the Bank of England has succeeded in achieving its goal of improving monetary policy transparency. All surprise measures show a clear downward trend in the frequency of policy surprises. Interest rates were raised at 19 meetings and lowered at 17 meetings. Of the 36 meetings at which the Bank rate was changed - 19 increases, 17 cuts - this policy action was expected by the market on 18 occasions, as measured by the Bloomberg survey. Of the remaining 18 instances, the market was either surprised that the MPC changed the policy rate or was surprised by the extent of the change. There were no instances where the market expected a change in the policy rate in the opposite direction to the change actually announced. Including the May 2000 meeting, at which the market expected a rate change but the MPC kept its repo rate constant, this adds up to a total number of 19 policy surprises according to the Bloomberg survey.[15] We therefore divide the meeting day sample into the 107 days when the change - including a change of zero basis points - in the policy rate was exactly as predicted - we term these 'No Unexpected Change Days' - and the 19 days when the rate changed by an amount different to market expectation - which we term 'Unexpected Change Days'.

[14]The period t policy announcement is classified as a surprise to the market if the difference between the period t+1 (IB or Liffe) rate and the period t-1 rate is greater than 10 (15) basis points.

[15]According to 'IB10' ('IB15', 'LIFFE10', 'LIFFE15'), 21 (13, 19, 10) policy surprises can be identified during the sample period.

4.5.1 Daily Data and Results

Daily observations of USD/GBP were obtained from the Federal Reserve Board. These are buying rates at noon New York time (17:00 London time). The daily data are sampled for the period May 1, 1997 to October 31, 2007.

Using daily data, we estimated the model represented by equation (4.1) above by OLS. The evidence indicates that the explanatory variables *Briefing*, *Day1*, *Day2*, and have no power in explaining exchange returns. This is true whether the meeting-related variables encompassed all meeting days or just those on which the Bank rate was changed unexpectedly. But the regression results do indicate the presence of significant GARCH effects. We then estimated the GARCH specification for daily exchange rate returns outlined above in which the dummy variables related to MPC meetings enter the conditional variance equation multiplicatively. Estimation is carried out by quasi-maximum-likelihood estimation using a Gaussian likelihood function and robust standard errors. The results indicate no explanatory power for variables related to all meetings. However, letting the meeting day dummies take the value one on meetings with surprising interest rate changes only, our preferred specification generates the values reported in Table 4.1. In this specification, the dummy for meeting day 2, *Day2* is statistically significant at the 1 percent level. The coefficient indicates that the conditional variance increases by about 25 percent on second meeting days. On briefing days, on the other hand, the conditional variance is about 11 percent lower than on non-meeting days.

Results regarding the *Day2* dummy were equivalent using all of our var-

Variable	Coefficient Estimate	p-value
Mean Equation		
Constant	1.13E-004	0.210
AR(1)	0.0358	0.069
AR(2)	-8.27E-003	0.680
AR(3)	-0.0429	0.039
Variance Equation		
Constant	3.81E-007	0.000
ARCH(1)	0.0286	0.000
GARCH(1)	0.955	0.000
Briefing	-1.10E-001	0.036
Day1	-4.32E-002	0.492
Day2	2.54E-001	0.004
Log Likelihood	10624.28	

Table 4.1: GARCH Model of Daily Exchange Rate Returns

ious alternative surprise measures based on 3-months short term interbank rates or 3-months sterling interest rate futures contracts (Tables A.5, A.6, A.7, A.8 in the Appendix), demonstrating the robustness of the results. So, although USD/GBP exchange rate returns appear to be unrelated to meeting day variables, exchange rate volatility is typically greater on days when policy surprises are announced.

4.5.2 Intraday Data and Results

The daily analysis indicates that second MPC meeting days are different from other days in volatility terms. We now take a microscopic look at these days in order to investigate systematic exchange rate movements and their determinants within a narrow window around the interest rate announcements. In this high-frequency setting, all references to MPC meeting days refer to second meeting days when the policy announcement is made. Tick data for USD/GBP were obtained from a major international bank for each of our 126 MPC meeting days and a set of 126 control days, defined as the same day of the week as the MPC exactly one week after the MPC meeting. Either no data or insufficient data could be made available to us for 14 out of the total of 252 days.[16]

We sample the last quotation of each 5-minute interval over the hours 7:00-17:00 London time to create a series of exchange rate returns, defined as the change in the logarithm of the 5-minute observations multiplied by 10,000.[17] The data for each day are stacked in serial order to create a data set with 28,556 observations. For further reference, it is important to notice that the 12:05 observation on any given day is the last quotation from within the interval 12:00-12:05.

The Markov model represented by equation (4.6) above was used to estimate the effect of MPC announcements on the transition probabilities. Estimation of the model was carried out using a modified version of the EM algorithm due to Diebold, Lee and Weinbach (1994). The two states are identified by significant shifts in the mean μ, the autocorrelation coefficicient ρ and the variance σ^2.[18] Recall that state 1 is the high-variance state associated with information-based trading and state 2 is the low-variance state associated with the normal market conditions of liquidity trading. The re-

[16]We also do not include the extraordinary and unscheduled meeting of September 18, 2001, and the respective control day.

[17]Danielsson and Payne (2002) compare one week of indicative quote data with firm quotes from an electronic FX brokerage and find that the properties of returns for each series become quite similar at a 5-minute sampling frequency. At higher frequencies, the indicative quotes tend to lag firm quotes. We choose the 5-minute sampling strategy to ensure that our exchange rate returns are representative of market conditions. The raw data were referenced to Greenwich Mean Time, so time references were appropriately adjusted to account for British Summer Time.

[18]Only in the case of the constant transition probability model are the means not significantly different from zero.

Coefficient Estimates (p-values)	
ρ(1)	-0.128 (0.000)
ρ(2)	-0.154 (0.000)
$\sigma^2(1)$	5.507 (0.000)
$\sigma^2(2)$	1.605 (0.000)
	Constant
P^{11}	1.68 (0.000)
P^{22}	1.74 (0.000)
LogL	-73022

Table 4.2: Constant Transition Probability Model

sults in Table 4.3 show that the estimated state 1 variance is generally found to be about 3.5 times that of state 2. Statistically significant negative first-order autocorrelation was also found in all models. Negative autocorrelation is a common finding in high frequency exchange rate returns.

In Table 4.2 we report estimates of the constant transition probability model and then in Table 4.3 we report the preferred model. The payoff from estimating the endogenous transition probabilities is demonstrated by the significant likelihood ratio statistic associated with comparing the constant transition probability model as the restricted estimate and the time-varying transition probability model as the unrestricted estimate.[19] In terms of the transition probabilities, P^{11} is the probability of remaining in the high-volatility state and P^{22} is the probability of remaining in the low-volatility state. Normally, we would expect $P^{22} > P^{11}$ and this is what the data reveal. Estimating a Markov-switching model with fixed transition probabilities re-

[19]From the log-likelihood values reported in Table 4.2 and 4.3, this statistic is -2(-73022+71082)=3880 (p-value = 0.00). Notice that the means are not significantly different from zero in the specification with constant transition probabilities.

Coefficient Estimates (p-values) of Regime Switching AR(1) Model

μ(1)	0.44 (0.000)	ρ(1)	-0.12 (0.000)	σ²(1)	6.11 (0.000)
μ(2)	-0.28 (0.000)	ρ(2)	-0.11 (0.000)	σ²(2)	1.64 (0.000)

Coefficient Estimates (p-values) for Endogenous Transition Probabilities

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	0.74 (0.000)	0.96 (0.000)	0.31 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.002)

LogL=-71082

Table 4.3: Preferred Time-Varying Transition Probability Model

sulted in the following estimates: P^{11}=(1.68)=0.95 and P^{22}=(1.74)=0.96. The unconditional probability of being in state 2 associated with these transition probabilities is given as $\frac{(1-P^{11})}{(1-P^{11})+(1-P^{22})} = 0.556$, so the unconditional probability of being in state 1 is 0.444.

Moving on to the time-varying probabilities model, it is first of all interesting to find that the mean return is significantly positive in the high-volatility state and significantly negative in the low-volatility state. This result might appear puzzling as it suggests that return volatility is high when the pound appreciates against the dollar and low when it depreciates. It is interesting to investigate, whether this result is driven by the interest rate announcement or whether it is simply an artefact of the data in the sample period considered. In order to further investigate this, we included additional intercept terms into the mean equation of our preferred specification as follows:

$$\Delta e_t = \mu(S_t) + \mu_{pos}(S_t) I(\Delta i > 0) + \mu_{neg}(S_t) I(\Delta i < 0) + \rho(S_t)[\Delta e_{t-1} - \mu(S_{t-1})] + \varepsilon_t \tag{4.8}$$

where μ_{pos} and μ_{neg} are additional constant terms, $I(\Delta i > 0)$ is an indicator function that takes the value 1 on interest rate surprise days between 12:05-13:00 if the announced interest rate is higher than expected, and $I(\Delta i < 0)$ is an indicator function that takes the value 1 on interest rate surprise days between 12:05-13:00 if the announced interest rate is lower than expected. We chose this definition for the indicator function as the analysis will proceed to show that the impact of interest rate announcements on the market is by far the greatest between 12:05-13:00 on announcement days when the announcement comes as a surprise to the market. As Table 4.4 reveals, the coefficients on μ_{pos} are significantly positive in both states and the coefficients on μ_{neg} are significantly negative in both states, with none of the previous results changed in a substantive way. These results indicate that, as expected, a higher UK policy rate than expected yields a positive mean return to holding sterling during the main impact period of the announcement, implying that the pound appreciates. A lower policy rate than expected yields a negative mean return during the main impact period of the announcement, implying that the pound depreciates. The finding that the mean return is, on average, generally positive in the high-volatility state and negative in the low-volatility state is therefore unrelated to the effect of the policy announcements. It rather appears to be a general artefact of the data within the sample period considered in this study: during periods unrelated

Coefficient Estimates (p-values) of Regime Switching AR(1) Model

μ(1)	0.43 (0.000)	ρ(1)	-0.13 (0.000)	$\sigma^2(1)$	6.09 (0.000)
μ(2)	-0.28 (0.000)	ρ(2)	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)
mupos(1)	3.27 (0.000)	muneg(1)	-0.98 (0.093)		
mupos(2)	0.53 (0.067)	muneg(2)	-2.54 (0.000)		

Coefficient Estimates (p-values) for Endogenous Transition Probabilities

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	0.75 (0.000)	0.86 (0.000)	0.31 (0.000)	-0.01 (0.000)	0.30 (0.031)	
P^{22}	7.24 (0.000)			-2.37 (0.000)		-0.89 (0.008)

LogL=-71064.25

Table 4.4: Markov Switching Model Including Additional Constant Terms

to surprising policy announcements, there is on average more volatility during times of appreciation than depreciation of the pound against the dollar.

Transition probabilities are modeled as varying with dummy variables that switch to 1 at certain times of day and are equal to 0 otherwise. Preliminary estimates suggested that the preferred model has P^{11} as a function of a constant and a dummy that is equal to one from 12:05-13:00 only on MPC meeting days when interest rates changed unexpectedly, a dummy equal to one from 12:05-13:00 on all MPC days, a dummy equal to one on all days between 12:05 and 13:45 and a dummy equal to one between 11:30 and 11:55 on all MPC days.[20] P^{22} is a function of a constant, a dummy equal to one on all days between 12:05 and 13:45, and a dummy equal to one on all MPC meeting days from 11:15-11:55.[21] Estimates are reported in Table 4.3, and

[20] Notice that the observation for 12:05 on any given day is the last quotation within the interval 12:00-12:05-i.e. the first observation in our data set after the interest-rate announcement.

[21] As discussed above, interest rates were deemed to have been changed unexpectedly by the MPC when the rate change, including a zero change, was different from the median expectation according to the Bloomberg survey of market participants - Tables A.2 to A.4.

indicate that each of the determinants of P^{11} and P^{22} differ significantly from zero with p-values of 0.01 or lower.

The results indicate that the probability of remaining in the informed trading state P^{11} is significantly higher from 12:05-13:00 following news that the MPC has unexpectedly changed its Bank interest rate. Based upon our preferred model specification, the estimated value of P^{11} changes from 0.74 before 11:30 to 0.85 between 11:30-11:55 and to 0.98 during the hour immediately following the unexpected change in the Bank rate.[22] The probability of remaining in the tranquil state, P^{22}, falls significantly between 11:15-11:55 on MPC announcement days and between 12:05-13:45 on all days. But although statistically significant, one may argue that the implied change in P^{22} is not economically significant. Tables A.9, A.10, A.11 and A.12 in the Appendix report results from estimating the same specification but replacing the Bloomberg survey surprise measure with the alternative measures detailed above. The similarity of results is striking and suggests that our findings are robust. Accordingly, for the remainder of our analysis we concentrate on the Bloomberg survey surprise measure.

Following on from our baseline estimates, Table 4.5 assesses the sensitivity of transition probability estimates over alternative specifications using afternoon dummy variables. In each case, the baseline model is augmented by an additional explanatory variable. These additional dummy variables are defined according to the same time divisions as previously, but over more types of day-second MPC days with and without interest rate changes, and all

[22]The preferred model specification is determined for surprises defined according to the Bloomberg Survey of market Economists.

Coefficient Estimates (p-values) for Additional Variables

	Additional 1	Additional 2	Additional 3	Additional 4	Additional 5
Type of Day	All	MPC	Surprise	All	Surprise
Time	12:05-13:00	12:05-13:45	12:05-13:45	11:30-11:55	11:30-11:45
P^{11}	-0.12 (0.260)	-0.16 (0.120)	0.09 (0.69)	0.02 (0.878)	0.37 (0.560)
P^{22}		-0.07 (0.782)	0.34 (0.688)		

	Additional 6	Additional 7
Type of Day	All	Surprise
Time	11:15-1155	11:15-11:55
P^{11}		
P^{22}	0.12 (0.770)	0.05 (0.964)

Table 4.5: Additional Variables Added to the Preferred Model Specification

days-than those incorporated in the preferred specification.[23] For instance, the dummy 'Additional 1' takes the value one from 12:05-13:00 on all days. Adding this dummy to the specification for P^{11} and testing its significance yields a coefficient of -0.12 and a p-value of 0.26. Table 4.5 indicates that none of the added variables is statistically significant.

One potential difficulty in this form of analysis is to ensure that estimated intraday state probabilities truly reflect the impact of MPC policy announcements, rather than the effect of other news or shocks. One obvious omitted variable candidate in this respect is the announcement calendar of other central banks. In particular, there are twenty-eight meetings in our sample where MPC announcement days coincided with policy announcements by the Governing Council of the European Central Bank (ECB). Announcements by the ECB occur at 12:45 GMT, which coincides with the reported significant increase in the probability of remaining in the informed trading state P^{11} following announcements of MPC policy decisions. To test whether sig-

[23] Recall that this focused only on 2nd MPC days with surprise announcements.

nificant volatility shifts in USD/GBP returns in part reflect a response to the publication of ECB interest rate decisions, we therefore included a set of dummy variables to proxy for these announcements. These dummies take the value one for time periods starting at 12:45 GMT on (a) all days on which MPC and ECB policy announcements coincided, (b) those coincident days on which the ECB announced an interest rate change, or (c) only those coincident days which involved an ECB policy surprise.[24] As Table 4.6 reports, only the dummy representing the time period 12:45-17:00 on all coinciding days was significant, when included in the specification for P^{11}, with a p-value of 0.02 and a coefficient of -0.21.[25] Overall, though, our results are not altered in any substantive way by the inclusion of any of these ECB dummy variables.

We can conclude that the evidence in presents a robust result: we have presented significant evidence of a systematic regime switch to a high-volatility informed trading state on MPC days when the BoE Bank rate is changed unexpectedly. This effect is highly significant for about an hour following the interest rate announcement. After this time, the probability of remaining in the informed trading state falls significantly. This result for MPC days with unexpected interest rate changes is clearly distinguished from other days and is not simply a "time of day" effect that exists in the market every day. In response to the questions posed above, Can we identify endogenous regime

[24]We define ECB surprise announcement days according to changes in the short term interbank rate (EURIBOR) using the same approach as for the BoE in Tables A.2 to A.4.

[25]A priori, one would expect the coefficient on this dummy variable to have a positive sign, indicating that ECB policy announcements increase the probability of higher return volatility in USD/GBP market. Accordingly, the observed negative estimated sign may indicate that this dummy is capturing something other than the volatility impact of ECB policy announcements.

Additional dummy variables added to the preferred model specification				
Equation	Day	Time	Coefficient (p-value)	LogL
P^{11}	ECB Coincides	12:45-14:00	-0.12 (0.255)	-71082
P^{11}	ECB Rate Change	12:45-14:00	-0.09 (0.636)	-71082
P^{11}	ECB Surprise Change	12:45-14:00	-0.54 (0.151)	-71081
P^{22}	ECB Coincides	12:45-14:00	-0.004 (0.992)	-71082
P^{22}	ECB Rate Change	12:45-14:00	-0.657 (0.402)	-71082
P^{22}	ECB Surprise Change	12:45-14:00	0.35 (0.777)	-71082
P^{11}	ECB Coincides	12:45-17:00	-0.21 (0.02)	-71078
P^{11}	ECB Rate Change	12:45-17:00	-0.06 (0.630)	-71082
P^{11}	ECB Surprise Change	12:45-17:00	-0.41 (0.056)	-71080
P^{22}	ECB Coincides	12:45-17:00	0.30 (0.182)	-71081
P^{22}	ECB Rate Change	12:45-17:00	-0.29 (0.473)	-71082
P^{22}	ECB Surprise Change	12:45-17:00	0.84 (0.330)	-71082

Table 4.6: Controlling for ECB Monetary Policy Announcements

switching? Are the transition probabilities driven by the news component in the policy announcements, we can answer with a strong affirmation.

We now turn to the final question to be addressed using intraday data: Is there evidence of positioning during MPC meetings prior to the policy announcement at noon on the second meeting day? The news anticipation effect is captured by the coefficients on the dummy variable for 11:30-11:55 on all MPC days in the P^{11} equation and the dummy variable for 11:15-11:55 on all MPC days in in the P^{22} equation.[26] As reported in Table 4.3, both dummy variables are indeed significant. The coefficients imply that from 11:30-11:55 there is an increase in the probability of remaining in the informed trading state-that is, state 1- and that from 11:15-11:55 there is a decrease in the probability of remaining in state 2, the liquidity trading state. The previous results summarized in Table 4.3 established that the noon announcement of

[26] Allowing the news anticipation effect to be captured by dummies with different starting points in the two equations improves the fit of the model.

unexpected interest rate changes were, indeed, price-relevant news as there is a switch to the high-volatility informed trading state immediately after the announcement. The current question requires that the pre-noon period receive a microscopic examination.

Parts A, B, and C of Table 4.7 incorporate alternative morning dummy variables into the preferred model as a further robustness check. This proceeds much like the analysis associated with the post-noon announcement effect. Starting with the baseline preferred model, we specify alternative dummy variables for the pre-noon period for our three different types of days: all days, all MPC meeting days, and MPC meeting days when an unexpected interest rate change was announced, and examine the sensitivity of the estimates to the additional variables. Part A includes dummy variables for all days over alternative times of the morning. For instance, the first row of part A includes a dummy equal to 1 from 11:45-11:55 in the P^{11} equation. The p-value indicates that this additional variable has no significant explanatory power. Our preferred model results are not altered by the inclusion of the variable. Similarly, the other variables added to the P^{11} and P^{22} equations have no significant explanatory power.

Part B of Table 4.7 incorporates additional morning dummy variables for all MPC days into the preferred model, and part C incorporates additional morning dummy variables on Unexpected Change Days. None of the dummies in part B are significant. By contrast, in part C dummies for the periods 11:45-11:55, 11:30-11:55 and 11:00-11:55 enter the P^{11} equation with significant p-values, and positive coefficients. In addition, the dummy for 9:00-11:55 in the P^{22} equation is borderline significant as well. This suggests

A. Alternative Specifications (morning dummies for all days)
Additional variables added to the preferred model specification

Equation	Variable	Coefficient (p-value)	Log Likelihood
P^{11}	11:45-11:55	-0.169 (0.300)	-71082
P^{11}	11:30-11:55	-0.08 (0.564)	-71082
P^{11}	11:15-11:55	-0.07 (0.500)	-71082
P^{11}	11:00-11:55	-0.000 (0.999)	-71082
P^{11}	9:00-11:55	-0.000 (0.999)	-71082
P^{22}	11:45-11:55	0.10 (0.862)	-71082
P^{22}	11:30-11:55	0.32 (0.459)	-71082
P^{22}	11:15-11:55	0.12 (0.769)	-71082
P^{22}	11:00-11:55	0.000 (0.999)	-71082
P^{22}	9:00-11:55	0.000 (0.999)	-71082

B. Alternative Specifications (morning dummies for all MPC days)
Additional variables added to the preferred model specification

Equation	Variable	Coefficient (p-value)	Log Likelihood
P^{11}	11:45-11:55	-0.02 (0.933)	-71082
P^{11}	11:15-11:55	-0.03 (0.879)	-71082
P^{11}	11:00-11:55	0.01 (0.930)	-71082
P^{11}	9:00-11:55	0.04 (0.530)	-71082
P^{22}	11:45-11:55	-0.44 (0.51)	-71082
P^{22}	11:30-11:55	0.53 (0.445)	-71082
P^{22}	11:00-11:55	0.66 (0.291)	-71082
P^{22}	9:00-11:55	0.18 (0.460)	-71082

C. Alternative Specifications (morning dummies for days with Surprise)
Additional variables added to the preferred model specification

Equation	Variable	Coefficient (p-value)	Log Likelihood
P^{11}	11:45-11:55	0.27 (0.047)	-71080
P^{11}	11:30-11:55	0.65 (0.038)	-71080
P^{11}	11:00-11:55	0.93 (0.041)	-71079
P^{11}	11:00-11:15	1.051 (0.161)	-71082
P^{11}	9:00-11:55	4.65 (0.953)	-71082
P^{22}	11:45-11:55	-0.28 (0.552)	-71082
P^{22}	11:30-11:55	-0.07 (0.935)	-71082
P^{22}	11:15-11:55	0.05 (0.964)	-71082
P^{22}	11:00-11:55	0.28 (0.828)	-71082
P^{22}	9:00-11:55	-1.88 (0.056)	-71082

Table 4.7: Markov-Switching Model of MPC News Anticipation Effects

that the probability of being in the high volatility state increases by to the interest rate announcement on days with unexpected interest rate changes as compared to days when the announcement is anticipated. This result could be interpreted as indicative of information leakages prior to the announcement. The results from our baseline specification are again not changed in a substantive way by the inclusion of these variables.

Taken as a whole, there is evidence of regime switching in terms of exchange rate volatility in the morning prior to the end of the MPC meetings. The evidence is strongest for the P^{11} equation for the 11:30-11:55 time period. During this interval, there is a statistically significant jump in the probability of remaining in the high-volatility state, from 0.74 to 0.85. Of course, since the meetings always end prior to the noon announcement and the MPC's policy decision is known by insiders, the regime switching could be a result of signals read by market participants. This is not to claim that there are deliberate information leaks emanating from the committee. It may be something much more subtle (recall the Greenspan briefcase story presented earlier). Furthermore, it may be that traders are simply closing down trade positions in order to limit their risk exposure precisely because they are unsure about the upcoming announcement. The evidence presented here indicates no particularly large probability shifts prior to the conclusion of MPC meetings. This is certainly true if one considers the probabilities of regime switching in the morning compared with the afternoon. The news impact of policy announcements appears to be much larger than any anticipation effect.

The implications of the intraday estimation results for the transition probabilities are summarized in Figure A.3 in the Appendix. The figure plots the smoothed unconditional probability of being in state 1, for the three types of days in our sample as generated by the preferred model reported in Table 4.3. This probability is averaged across all observations for each type of day for each 5-minute interval. One can observe dramatic differences across types of days and time of day.

It is clear that non-MPC meeting days are characterized by low-volatility, liquidity trading as the probability of remaining in the informed trading state is quite low all throughout the day; fluctuating between 0.25 and 0.45. On MPC meeting days when no unexpected interest rate change occurs, there is an increase in the average unconditional probability of being in state 1 that begins modestly around 11:30 and continues until 12:05 when it jumps to about 0.53. After this peak, the probability quickly falls to about 0.40 by about 12:30 and then by 13:00 is quite similar to the afternoon pattern on non-MPC days.

On MPC meeting days when an unexpected interest rate change occurs, however, there is a dramatic jump at noon when the policy announcement is released, from about 0.55 to more than 0.90. The probability of being in the informed trading state subsequently remains above 0.70 until about 13:00 after which it continues to fall so that by about 13:30 it appears to follow a pattern much like other days. In Figure A.4 in the Appendix we have plotted the same information as in Figure A.3 but using alternative measures of policy surprises, as detailed above. The overall pattern is strikingly similar, suggesting that the results are robust to the exact characterization of policy surprises.

An interesting feature of Figure A.3 is that the probability of moving into the high-volatility state rises even on days when the interest rate decision was correctly anticipated. This is perhaps worthy of further investigation, since one might expect anticipated announcements to be discounted into the exchange rate prior to the announcement. One possible explanation of this finding may be related top the fact that we have used the median expectation from the Bloomberg survey and ignored any dispersion in expectations among survey participants: there will in general still be some people surveyed who are surprised by the announcement even when it coincides with the median view. If these people then initiate trades in response, this may then generate a series of "hot-potato" trades, although the multiplier effect on trades would be expected to be smaller than if the majority of the market were surprised.

In order to investigate the validity of this argument, we use the measure of forecast dispersion introduced in Tables A.2 to A.4 to distinguish days when analysts were unanimous in regarding their expectation of the policy announcement from MPC days when they were not. We construct a dummy that takes a value one from 12:05-13:00 on MPC days without policy surprises when the standard deviation of analysts' expectations is zero. We then construct a second dummy variable that takes the value one from 12:05-13:00 on MPC days without policy surprises on which the standard deviation is greater than zero. We include both of these variables in the P^{11} equation of our preferred specification and exclude 'Dummy 2', representing the time span 12:05-13:00 on all MPC days. The data set is reduced to 24,683 observations due to a lack of data on analysts' expectations prior to October 1998.

Coefficient Estimates (p-values) of Regime Switching AR(1) Model

μ(1)	0.38 (0.009)	ρ(1)	-0.13 (0.000)	σ²(1)	5.90 (0.000)
μ(2)	-0.25 (0.000)	ρ(2)	-0.11 (0.000)	σ²(2)	1.55 (0.000)

The number of observations is reduced to 24683 for this regression due to lack of data on Individual analysts' expectations prior to October 1998.

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	1.14 (0.000)	1.14 (0.000)	Excluded	-0.01 (0.000)	0.17 (0.240)	
P^{22}	7.20 (0.000)			-2.45 (0.000)		-1.02 (0.003)

	Dummy 6	Dummy 7
Type of Day	MPC	MPC
Surprise	No	No
Dispersion	Yes	No
Time	12:05-13:00	12:05-13:00
P^{11}	0.40 (0.001)	-0.12 (0.478)

LogL=-59477

Table 4.8: Markov-Switching Model Taking Account of Forecast Dispersion

Table 4.8 shows that only the variable indicating non-surprise MPC days with dispersion in expectations exhibits a significant influence on P^{11}, with a p-value of 0.001 and a coefficient of 0.4. This suggests that the rise in P^{11} at noon on MPC days without policy surprises is only significant when at least one individual deviated from the median market expectation. This finding gives support to the argument above that the increase in P^{11} at noon on MPC days without surprises is due to the use of the median individual analyst's expectation as a proxy for the entire market's expectation. In other words, even if the median expectation does not differ from the actual announcement, there are still market participants who are surprised by the announcement. This emphasizes the importance of investor heterogeneity in the foreign exchange market (Sager and Taylor, 2006).

Overall, the evidence in Figure A.3 indicates that MPC days are, indeed, different from other days. The noon policy announcement appears to be price-relevant news, in particular when the announcement comes as a surprise to the market. There is some modest evidence of positioning in advance of the announcement on all MPC days, but for days when interest rates are changed unexpectedly, it appears that the market response comes immediately at noon with the news. It also appears that the market takes around an hour to digest the news component of an unanticipated announcement in terms of the average P^{11} dropping significantly back to around its previous level.

It should also be noted that our findings-in particular the evidence of a strong exchange rate reaction to the news announcement (which is much more marked on days when the interest rate announcement differs from the median market expectation) with little strong evidence of positioning during the morning period of the meeting-are qualitatively similar to those reported by Sager and Taylor (2004) in their high-frequency study of the exchange rate effects of interest-rate announcements by the Governing Council of the ECB, suggesting that the results are robust.[27]

It is also interesting to contrast our results with those of Evans and Lyons (2007). We conclude in favour of a significant but relatively short-lived impact upon the volatility of exchange rate returns for both unexpected and expected rate changes. By contrast, Evans and Lyons' analysis of proprietary

[27]Likewise, Clare and Courtenay (2001) examine the response of interest and exchange rates to UK monetary policy announcements and macroeconomic data releases using 1-minute tick data in a sample that spans the introduction of operational independence at the BoE. They, too, find in favour of a significant volatility effect due to both types of new information, and for both interest and exchange rates, and also conclude that the implications of policy innovations are more quickly incorporated into interest and exchange rates in the post-independence era than previously was the case.

order flow data concludes in favor of a very persistent relationship between order flow and exchange rate returns, with the former exhibiting out-of-sample predictive power for returns as much as one quarter ahead, but no significant impact in the short-term. This contrast reflects differences in the behaviour of market participants in the various segments of the foreign exchange market. In this paper, we have isolated the impact of knee-jerk trading on the volatility of returns around the time of MPC interest rate announcements, as inter-dealer positioning adjusts to reflect the arrival of this new information. This is an important and quick process, as befits a liquid and relatively efficient market as foreign exchange. But it is only part of the story. Evans and Lyons (2007) focus explicitly away from inter-dealers and on the customer segment of the market that accounts for more than 50 percent of market turnover.[28] As Sager and Taylor (2006) discuss, other than smaller hedge funds the majority of foreign exchange market customers typically does not a similar exhibit knee-jerk reaction to news as the inter-dealer market. Although this behavior contradicts the Rational Expectations Hypothesis, it is rational - in the sense of being profit-maximizing - and reflects both the size of assets under management, and associated transaction costs of trading, and that a large proportion of the trading activity of this market segment is not driven by news innovations, but benchmark adjustments (Lyons, 2001).

[28]This segment includes asset management firms, such as mutual fund managers, as well as hedge funds, corporates and central banks. For information on the share in foreign exchange market turnover of the various market segments, see BIS (2007).

4.6 Discussion

Following the granting of independence on the setting of interest rates to the Bank of England in 1997, the Monetary Policy Committee was created as its interest-rate setting committee, charged with fostering monetary policy consistent with stable inflation and economic growth. A stated aim of the new policy regime was that monetary policy should be more transparent than hitherto. The availability of MPC decisions affords us a rare opportunity to examine how the decisions of the key policy-setting committee are impounded into the foreign exchange market.

Since the MPC meets at regularly scheduled, pre-announced times and the policy decision is always announced at noon, the meetings provide a natural laboratory for examining exchange rate dynamics on days when monetary policy is formulated and announced. Our particular interest is with respect to the news content of the policy announcement and also whether there is any evidence of positioning in the foreign exchange market during the MPC meeting prior to the announcement.

We employed daily data on USD/GBP to analyze any differences that may exist in the behavior of exchange rate returns on the three kinds of days associated with MPC meetings: the pre-meeting briefing day; the first day of the meeting; and the second day of the meeting when the policy announcement is made. We estimated models of daily exchange rate returns to infer if information on MPC meeting days contains any explanatory power. Our estimation results suggest that daily exchange rate returns are well characterized by mean-zero changes and meeting day information has no explanatory power for returns. But modeling the conditional volatility of the daily returns revealed evidence of significantly greater volatility on second meeting days when interest rates are changed unexpectedly.

Given this result, we turned to a microscopic view of second meeting days using intraday exchange rate returns and an endogenous-probability Markov-switching framework. Our estimated model assumed that there exist two states: state 1, the high-volatility state associated with informed trading, and state 2, the low-volatility state associated with liquidity trading. We diverged from the usual non-linear regime-switching framework to model endogenous transition probabilities as a function of information regarding the meeting days. The transition probabilities were found to switch systematically and significantly on meeting days. The probability of remaining in the high volatility state was estimated to increase from 0.74 before 11:30 to 0.98 from 12:05-13:00 on MPC meeting days when interest rates are changed by an amount different from that expected by the market (or are not changed when the market expects a change).

The second day of MPC meetings, the day on which interest rate decisions are announced, is therefore best characterized as having a-statistically and economically-significant exchange rate reaction to the news announcement at noon with some evidence of positioning during the morning period of the meeting. These announcement effects last for around an hour to ninety minutes and are much more marked on days when the interest rate announcement differs from the ex ante median market expectation.

An interesting extension of these results would be to empirically test the ability of market participants to profitably exploit these announcement effects-that is, to validate the economic significance of our findings-through a profit-loss analysis of trading strategies that, say, introduce short-lived option structures in USD/GBP on the second day of MPC meetings around the time of the policy announcement. This is a task we leave to future research.

Chapter 5

Monetary Shocks and Portfolio Choice

5.1 Introduction

The current financial crisis has for several years been preceded by substantial global imbalances in trade and capital flows.[1] In particular the United States were not only at the center of the financial crisis, but also among the economies relying most heavily on capital inflows to finance a growing trade deficit. A number of observers have argued that accommodative monetary policy over the past decade has been a key culprit behind these imbalances by inducing the build-up of excess liquidity, a rise in financial leverage and a boom in asset prices. This, in turn, may have contributed to a surge in private consumption, in part due to wealth effects, and ultimately a rising US current account deficit (e.g. Taylor 2009).

[1]This chapter is co-authored with Marcel Fratzscher and Roland Straub

During the same period, capital flows to the United States have exhibited peculiar dynamics regarding their composition, with the US current account deficit financed increasingly via inflows into bonds as opposed to equities. Figure 1 illustrates this point, underlining that in particular since 2001, in an environment of accommodative monetary policy, net inflows into US debt securities have surged to close to 6% of US GDP or about USD 800 billion per year, while net inflows into equities, FDI and other investment have been modest and even negative at times.

The role of monetary policy thus warrants closer scrutiny in order to understand how it may have contributed to the dynamics of capital flows, both in terms of their size and their composition. This is a first objective of the paper. More specifically, the paper focuses on the effect of monetary policy shocks in the United States on the US trade balance and different types of capital flows.

Moreover, the focus on the effect of monetary policy shocks on the direction and composition of capital flows allows us to contribute to the debate on the determinants of portfolio choice, and how asset price movements are related to portfolio decisions of investors across countries as well as across financial asset classes. This is the second objective of the paper. An important strand of this literature analyses portfolio rebalancing versus return chasing as motives for investment decisions, in an environment of incomplete financial markets and imperfect substitutability of financial assets (e.g. Bohn and Tesar 1996, Hau and Rey 2006 and 2008, Albuquerque 2007, Devereux and Sutherland 2006, Tille and van Wincoop 2007). A related literature focuses on understanding asset price comovements, in particular the peculiar stock-bond return correlation, which are hard to explain via the use of empirical models to date (e.g. Shiller and Beltratti 1992, Baele, Bekaert and Inghelbrecht 2008).

The paper employs a standard structural VAR specification to identify monetary policy shocks, relying on sign restrictions imposed on the impulse response functions of a few macroeconomic variables, following closely Canova and De Nicolo (2002), Uhlig (2005) and Fratzscher and Straub (2008). We specify our Bayesian VAR using US variables relative to those of other G7 members in the baseline specification, and relative to an extended sample of rest of the world countries in the robustness specification.

The empirical analysis yields two key findings. First, US monetary policy shocks exert a statistically and economically meaningful effect on US capital flows and the trade balance. An exogenous easing of US monetary policy by 100 basis points (b.p.) induces net capital inflows and a worsening of the US trade balance of around 1% of GDP after 8 quarters. The variance decomposition indicates that US monetary policy shocks over the period 1974 to 2007 explain about 20-25% of the variation in both the US trade balance and capital flows at that horizon. As to the channels, it appears that wealth effects play a central role. Equity returns rise on impact by about 6% in response to a 100 b.p. policy easing,[2] while interest rates fall. Both of these responses in turn induce an increase in private consumption for about 8 quarters, and thus a deterioration in the trade balance.

The second main finding regards the effect of monetary policy shocks on

[2]This estimate is essentially the same as that found in the literature (Rigobon and Sack 2002, Ehrmann and Fratzscher 2004, Bernanke and Kuttner 2005) which mostly use an event-study methodology focusing on the daily response to FOMC policy surprises.

the composition of US capital flows. The intriguing finding is that an exogenous US monetary policy easing causes *net inflows* into debt securities, foreign direct investment (FDI) and other investment, while inducing *net outflows* of portfolio equities. Monetary policy shocks thus entail a conditional negative correlation between flows of portfolio equity and debt. By contrast, monetary policy shocks induce a positive conditional correlation in equity returns and bond returns, as is well known in the literature. Moreover, they cause a negative conditional correlation between equity flows and equity returns, but a positive conditional correlation between bond flows and bond returns. The findings are robust to a battery of extensions and sensitivity checks, such as using the approach suggested by Fry and Pagan (2007) to extract the median impulse responses from a single model.

How should one understand and rationalize these empirical findings? From an observational perspective, the findings seem to fit well with the stylized facts of Figure 1 stressing the shift in the composition of capital inflows from equities to bonds amid an environment of low interest rates in recent years.

Furthermore, our empirical analysis allows us to contribute to the literatures on the determinants of portfolio choice as well as on asset price comovements. As to the literature on portfolio choice, one strand of the debate has emphasized the role of a return chasing motive behind international capital flows, in which investment decisions are primarily driven by expected returns. Bohn and Tesar (1996) analyses return chasing and portfolio rebalancing in a simple ICAPM framework, yielding a decomposition of net purchases into transactions necessary to maintain a balanced portfolio and net purchases that are triggered by time varying investment opportunities. Their results suggest that US transactions in foreign equities are primarily driven by the latter return chasing motive. In another classic paper, Brennan and Cao (1997) study the effect of information asymmetries between domestic and foreign investors on international portfolio flows, finding evidence in favor of a positive correlation between equity flows and returns, though only for US investments abroad.[3] Finally, an important recent strand of this literature rationalists the return chasing motive for capital flows (Albuquerque 2007), and presents evidence that such return chasing is taking place at a global scale due to asymmetries in information and differences in investor performance (Albuquerque et al. 2008).

Another strand of the literature has provided evidence in favor of a prominent role for the *portfolio rebalancing motive* as a driver of capital flows (Branson and Henderson, 1985). The more recent literature stresses the incompleteness of financial markets and the role of various forms of risk that make domestic and foreign assets imperfect substitutes, and in which thus *portfolio rebalancing* is a key driver of international capital flows. Hau and Rey (2006) argue that the (unconditional) negative correlation between equity returns and exchange rate returns may be rationalized through a portfolio rebalancing motive in which exchange rate risk induces investors to reallocate capital out of countries with rising exchange rates.

In Hau and Rey (2008), the authors find evidence in favor of portfolio rebalancing in a sample of 6500 international equity funds for the US, UK,

[3]However, they do not find such a positive correlation for foreign investment into US assets. They explain this finding by the notion that foreigners are less informed and thus revise their predictions more strongly when they receive a given signal.

Canada and the EU. Calvet et al. (2009) find similar micro evidence for portfolio rebalancing in the behavior of Swedish households. Froot and Ramadorai (2002) use proprietory data on daily institutional investor currency flows and find that these flows are highly correlated with contemporaneous and lagged exchange rate changes. At a macro level, there is a growing literature emphasizing the role of country risk and market frictions for why capital is not flowing to countries with high asset returns (e.g. Kraay et al. 2005, Gourinchas and Rey 2006, Gourinchas and Jeanne 2006, Lane and Milesi-Ferretti 2006, Daude and Fratzscher 2007, Alfaro et al. 2008).[4]

The evidence of the present paper is consistent with a portfolio rebalancing motive for equity portfolio flows - as implied in the negative conditional correlation between equity returns and equity flows - and a motive akin to return chasing behind investment decisions for bonds - as indicated by the positive conditional correlation between bond returns and bond flows.

As regards the return correlations, the literature on asset price comovements stresses that there tend to be strong time variations in the comovements of returns across different asset classes, such as between equity returns and bond returns. These strong time variations constitute a puzzle, as neither present value models (Shiller and Beltratti 1992), nor consumption-based asset pricing models (Bekaert, Engstrom and Grenadier 2005), nor dynamic factor models with a broad set of economic state variables (Baele, Bekaert and Inghelbrecht 2008) are able to explain them well. Andersen et

[4]An important related literature is emerging using DSGE models with endogenous portfolio choice (Coeurdacier 2005, Devereux and Sutherland 2006, Tille and van Wincoop 2007, Pavlova and Rigobon 2008), which stresses the imperfect tradability of risk, which contributes to the home bias in the international investment patterns.

al. (2007) show that the bond-stock return correlation is positive during periods of expansion but negative and large during economic contractions. They conjecture that this strong time variation and switch in sign in the correlation may be explained by the time-variation in the relative importance of cash flow effects and discount rate effects: during expansions, discount effects dominate thus inducing a positive correlation between stock and bond returns; while cash flow effects are dominant in contractions so that returns on bonds - with fixed nominal cash flows - have the opposite sign compared to returns on equities - which have stochastic dividends.

The present paper stresses that this positive correlation between stock returns and bond returns is present precisely when discounting effects (monetary policy shocks) dominate. Of course, it also implies that this correlation may be different when other shocks dominate. As such, the present paper focuses on understanding the effect of one specific shock for portfolio choice and asset prices, while we leave it for future research to condition the analysis on other types of economic shocks. Moreover, the paper's findings emphasis the importance of jointly analyzing quantities and prices, i.e. portfolio flows in conjunction with asset price movements, and also across asset classes for understanding the portfolio choices of investors.

The paper proceeds as follows. In Section 2, we examine the determinants of net capital flows in a simple intertemporal capital-asset pricing model as discussed in Bohn and Tesar (1995). Section 3 presents the empirical model and outlines methodology used to identify monetary policy shocks in detail. Section 4 presents the empirical findings for the benchmark specification and discusses the interpretation and the implications of the results. Robustness and sensitivity tests are presented in section 5. Section 6 concludes.

5.2 Decomposing Net Capital Flows

We begin by examining the determinants of net capital flows in an intertemporal capital-asset-pricing model as discussed in Bohn and Tesar (1995). We use the model to fix language and notation. Although originally constructed for equity investment, the intutition of the model can be applied to most other forms of investment in a similar fashion. The model yields a natural decomposition of net purchases of assets into (i) transactions that are necessary to maintain a balanced portfolio of securities (*portfolio-rebalancing* effect) and (ii) net purchases that are triggered by time-varying investment opportunities (*return-chasing* effect). As a result, depending on which of the two effects dominate in the investor's portfolio allocation, the correlation between investment returns and net capital flows may take either sign.

We begin by considering the problem an investor faces who can purchase both domestic and foreign equity. Let NP_{kt} be the period t net purchases of stocks in country k and let W be the value of the investor's portfolio. By definition, the following relationship determines how net purchases of asset k are related to portfolio shares (α_{kt}) and total wealth (W_t):

$$NP_{kt} = \alpha_{kt} W_t - (1 + g_{kt})(\alpha_{kt-1} W_{t-1}) \tag{5.1}$$

where g_{kt} is the capital gain on security k. Since wealth at time t is a function of the return on the *total* portfolio between periods $t-1$ and t, net purchases can be approximated as:

$$NP_{kt} = (\alpha_{kt} - \alpha_{kt-1})\, W_{t-1} + (d_t^p + g_t^p - g_{kt})(\alpha_{kt-1} W_{t-1}) \qquad (5.2)$$

where d_t^p and g_t^p are the dividends and capital gains on the investor's total portfolio. The right hand side of equation (5.2) comprises two terms, each representing possible motives for the investor to purchase or sell security k. The first component indicates that a change in the investor's desired portfolio weight on security k between period $t-1$ and t may trigger the purchase or sale. The second component suggests that the investor will purchase security k when her wealth increases due to dividend payments on her total portfolio of assets. However, she will sell security k when returns on asset k exceed returns on the rest of the portfolio such that the portfolio is not in balance anymore.

In particular, Cox et al. (1985) show that investors facing the standard trade-off between mean return and variability behave subject to the following optimality condition linking the portfolio weight on security k to the return process:

$$\alpha_{kt} = \sigma e_k \Sigma_t^{-1} E_t(\mu) + \eta_{kt} \qquad (5.3)$$

where σ is the coefficient of relative risk aversion, $E(\mu_t)$ is the vector of expected excess returns on all securities, e_k is a 0-1 vector that selects element k, Σ_t is the covariance matrix of returns and η_{kt} is the component of the portfolio used to hedge the investor against all other types of risks that are not related to her equity investment strategy. For tractability, it is assumed that time variation in the model occurs in the first moments of the driving processes for returns and the state variables only. The reason is that this ensures that the hedge terms are constant. Substituting into (5.2) yields a condition that determines the investor's portfolio adjustment strategy:

$$NP_{kt} = (d_t^p + g_t^p - g_{kt})(\alpha_{kt-1}W_{t-1}) + \sigma e_k \Sigma_t^{-1} \left[E_t(\mu) - E_{t-1}(\mu)\right] W_{t-1} \quad (5.4)$$

The first component on the right hand side of equation 5.4 captures what we denote the *portfolio-rebalancing* effect, namely net purchases of asset k that are required to maintain constant portfolio weights. The second term captures the extent to which investors adjust portfolio weights as the portfolio is reoptimized over time. Given a fixed level of risk aversion and a constant variance-covariance matrix of returns, an investor adjusts portfolio weights only if his expectations of excess returns are revised over time. We therefore refer to this as the *return-chasing* effect.

The two effects imply different correlation structures between the (expected) return on capital and capital flows. If the *portfolio rebalancing* effect dominates, an increase in the relative return on assets in country k should lead to a net capital outflow as indicated by the negative coefficient on the local capital gain g_{kt}. On the other hand, if the *return chasing* effect dominates then changes in the investor's expectation of excess returns in country k should dominate portfolio flows. The latter implies a positive correlation between expected excess returns and net capital flows.

We emphasise that the purpose of this section is purely motivational in order to illustrate the implications of changes in returns for portfolio flows, and vice versa. Our empirical exercise in the next sections will investigate which effect dominates empirically when analysing net portfolio flows of debt and equity following a monetary policy shock in the United States.

5.3 The Empirical Model

In this section, we present our empirical model and explain the implementation of our pure-sign restrictions approach. In Appendix A.13, we define further the variables that we use in the analysis and declare the respective data sources.

5.3.1 Model Specification

We estimate a structural VAR model of the form

$$y_t = c + \sum_{i=1}^{p} A_i y_{t-i} + B^{-1}\varepsilon_t \tag{5.5}$$

where B is an $(n \times n)$ matrix of contemporaneous coefficients, A_i is an $(n \times n)$ matrix of autoregressive coefficients, ε_t is an $(n \times 1)$ vector of structural disturbances and y_t an $(n \times 1)$ vector of endogenous variables, and p is the number of lags in the VAR. The model we use is of dimension $n = 8$, where y_t is defined as

$$y_t = [c_t - c_t^*, i_t - i_t^*, cpi_t - cpi_t^*, eq_t - eq_t^*, nb_t, reer_t, tb_t, cap_t] \quad (5.6)$$

The variable cap_t represents the different capital flow variables that are included in the model one at a time. These are the aggregate *Financial Account*, and its four individual components *Foreign Direct Investment*, *Portfolio Equity, Portfolio Debt* and *Other Investment.* All of these are net flows, i.e. changes in assets minus liabilities, and are by definition relative variables where a positive value denotes a net inflow of the respective type of capital from the rest of the world into the United States.[5] The remaining variables are the trade balance as a ratio of GDP, tb_t, which is the main counterpart of the financial account in the balance of payments identity.[6] We include this variable not only because at any point in time trade deficits must be balanced by financial account surpluses, but also because potentially net financing needs for trade deficits may be an important driving factor of capital flows.[7]

The variable nb_t is the ratio of non-borrowed to total reserves which we include solely for the purpose of identifying monetary policy shocks. We follow Strongin (1995), Faust and Rogers (2003) and Uhlig (2005), among others, who argue that the reserve ratio is the monetary aggregate that can be most closely associated with changes in the monetary policy stance of the

[5]A problem associated with the capital flow variables is that they tend to be very volatile at times. We therfore use five-quarter moving averages of them for our estimation.

[6]More precisely, given the accounting identity of the balance of payments, the financial account tracks the trade balance quite closely (with the opposite sign) with differences arising due to the income and transfer accounts under the current account, the capital account, changes in reserves and statistical discrepancies.

[7]The qualitative results of this study are robust to the inclusion of the current account balance in place of the trade balance in the model.

United States. The variable $reer_t$ is the log of the real effective exchange rate. The remaining variables $c_t - c_t^*, i_t - i_t^*, cpi_t - cpi_t^*$ and $eq_t - eq_t^*$ respectively represent the percentage differences between US and rest of the world variables for consumption ($c_t - c_t^*$, in USD), short term interest rates ($i_t - i_t^*$), CPI inflation ($cpi_t - cpi_t^*$) and equity returns ($eq_t - eq_t^*$, in USD).

In the benchmark specification, "Rest of the World" is defined as a GDP weighted average of the non-US G7 countries. The exception is the equity return differential $eq_t - eq_t^*$ for which countries are given weights according to their share in the aggregate non-US G7 equity market capitalization. It is important to note that the given definition of the equity return differential $eq_t - eq_t^*$ in USD terms implies that it effectively represents the deviation from equity parity as defined in Hau and Rey (2006). According to their definition, equity parity holds whenever the equity return differential in local currency terms is exactly offset by the nominal exchange rate return.

We use quarterly data that spans the sample period between 1974 and 2007. The year 1974 is used as a starting point of the analysis as it marks the beginning of the floating exchange rate period after the collapse of the Bretton Woods system. The choice of the time period and the associated data availability considerations naturally limit the choice of candidate countries to be included in our definition of the "Rest of the World". We will, however, test the robustness of our results to the extension of the definition by including a range of additional countries. The economies included in the two "Rest of the World" samples are listed in the Appendix.

5.3.2 Identification of Monetary Policy Shocks

We are interested in the effect of a monetary policy shock on the different types of net capital flows between the United States and the rest of the world. We identify monetary policy shocks using the pure-sign restrictions approach pioneered by Faust (1998), Canova and de Nicoló (2002) and Uhlig (2005). The technique allows us to identify structural error terms from a reduced form version of the VAR model presented in equation (5.5) by using a minimum of intuitively appealing sign restrictions on the impulse response functions of some of the endogenous variables included in the vector y_t. The identification restrictions we use are well grounded in economic theory and are by now widely used in the academic literature to identify monetary policy shocks.

We present the restrictions we use to identify an expansionary monetary policy shock in Table 5.1. An upward arrow indicates that the respective variable is required to increase for four quarters following the shock. In particular, we assume that an expansionary monetary policy shock reduces short term interest rates and has a positive effect on consumption, inflation and the ratio of non-borrowed to total reserves. In terms of the relative variables in our model, this implies that a monetary policy shock reduces $i_t - i_t^*$ and increases $c_t - c_t^*$, $cpi_t - cpi_t^*$ and nb_t. In Table 5.1, the upward arrow on $c_t - c_t^*$ is shown in parentheses as we will leave out this restriction at a later point in the analysis as a robustness check. Table 5.1 also presents the restrictions of two additional types of shocks: an aggregate demand shock and an aggregate supply shock. The reason is that it has been shown that

Table 5.1: Sign Restrictions

Structural Shock	$i_t - i_t^*$	$cpi_t - cpi_t^*$	nb_t	$c_t - c_t^*$
Monetary Policy	$\uparrow$	$\uparrow$	$\uparrow$	$(\uparrow)$
Aggregate Demand	$\uparrow$	$\uparrow$		$\uparrow$
Aggregate Supply		$\downarrow$		$\uparrow$

increasing the number of identified shocks can help to uncover the correct sign of the impulse response functions (Paustian 2007). We therefore identify these two additional structural shocks as a robustness check but do not report results on the impulse responses to these shocks in what follows. Moreover, the table illustrates that the identifying restriction for monetary policy shocks make this type of shock distinct from supply and demand shocks.

We now move to the implementation of the pure-sign restrictions approach. Thereby, we follow Canova and De Nicoló (2002), Uhlig (2005) and Peersman and Straub (2008) in recovering the structural error terms from the estimated reduced form model via the use of sign restrictions on the impulse response functions of some of the endogenous variables. Let us first define $v_t = B^{-1}\varepsilon_t$ as the reduced form residuals of the VAR. Standard OLS estimation of the reduced form VAR yields thereby an estimate of the variance-covariance matrix $\Sigma = E\left(v_t v_t'\right)$. In order to identify the structural error terms underlying these disturbances such that impulse response functions can be constructed, we need to find a way to choose among the infinite number of possible decompositions of Σ. Two candidate decompositions are the Cholesky Factor and the eigenvalue-eigenvector decomposition. The latter is given by $\Sigma = CC' = PDP'$ where P is a matrix of eigenvectors and D is a diagonal matrix with the corresponding eigenvalues on the main diagonal. Although this decomposition of the variance-covariance matrix is intuitively not very appealing as it is economically not meaningful, the crucial advantage is its uniqueness and the fact that it generates orthonormal shocks. It therefore allows us to generate any possible decomposition of Σ by finding an orthonormal matrix Q such that $QQ' = I_n$ and writing the newly found candidate decomposition as $\Sigma = CQQ'C' = \widehat{C}\widehat{C}'$.

The only task left is thus to find an algorithm that allows to efficiently search through the infinite space of orthonormal matrices Q and to construct candidate decompositions accordingly. Such an algorithm can be achieved by constructing a desired number of orthonormal matrices Q as $Q = \prod_{m,n} Q_{m,n}(\theta)$ where $Q_{m,n}(\theta)$ are rotation matrices of the form:

$$Q_{m,n}(\theta) = \begin{bmatrix} 1 & \cdots & 0 & \cdots & 0 & \cdots & 0 \\ \cdots & \ddots & \cdots & \cdots & \cdots & \cdots & \cdots \\ 0 & \cdots & \cos(\theta) & ... & -\sin(\theta) & \cdots & 0 \\ \vdots & \vdots & \vdots & 1 & \vdots & \vdots & \vdots \\ 0 & \cdots & \sin(\theta) & \cdots & \cos(\theta) & \cdots & 0 \\ \cdots & \cdots & \cdots & \cdots & \cdots & \ddots & \cdots \\ 0 & \cdots & 0 & \cdots & 0 & \cdots & 1 \end{bmatrix} \tag{5.7}$$

and where m and n are the rows that are being rotated by the angle θ. The number of rotation matrices is naturally large in a model of dimension $n = 8$. In fact, there are $n(n-1)/2 = 28$ bivariate rotations for a given angle θ. The algorithm we use entails to randomly draw a rotation angle $\theta\,(j)$ and to construct the matrix Q as the product of the resulting rotation matrices $Q_{m,n}(\theta)$. In principle, any rotation can be constructed by varying the parameter $\theta\,(j)$ in the range $[0,\pi]$. This algorithm allows for an efficient exploration of the infinite space of possible realizations of the matrix Q. A given draw Q_j allows us to construct a candidate contemporaneous impact multiplier matrix $\widehat{C_j}$ and the corresponding set of impulse response functions

$$R_{j,t+k} = A(L)^{-1}\widehat{C}_j v_t \tag{5.8}$$

For estimation and inference, we use a Bayesian approach. Our prior and posterior for the coefficient matrix $A\,(L)$ and the variance covariance matrix Σ come from the Normal-Wishart family. We use the same weak parameterization for the prior as in Uhlig (2005). In order to draw "candidate truths" $\widehat{C_j}$, we jointly draw from the Normal-Wishart posterior for the variance covariance matrix Σ_j and the coefficient matrix $A_j(L)$ as well as from the uniform distribution of the rotation angles $\theta\,(j)$. Impulse response functions are then constructed using the above procedure. A sign restriction on the impulse response of variable p to shock q after k periods following the shock is of the form:

$$R_k^{pq}\,(j) \gtrless 0 \tag{5.9}$$

We impose our set of sign restrictions for four periods after the shock occurs. If the impulse response functions obey the postulated sign restrictions, we keep the draw. If they do not, we discard it. We continue this procedure until we have found 1000 admissible draws. When identifying multiple shocks, we identify them simultaneously. This means that, in order for a given draw to be accepted, it must obey the restrictions applicable to each of the identified shocks simultaneously.

At any point in the response horizon the distribution of impulse responses across accepted draws is subject to two different sources of variation. One is the uncertainty around the estimates of $A(L)$ and Σ which we take into account by means of using a Bayesian approach. The other source of variation is the uncertainty introduced by drawing a value for the candidate rotation angle $\theta(j)$. This effectively generates variation "across models" and brings about the question of how to summarize the distribution of impulse response functions across accepted draws at each point along the response horizon. Following most of the sign restrictions literature, we simply report the median of each distribution along with upper and and lower quantiles in order to give an idea of the range of permissible impulse responses. Fry and Pagan (2007) criticize this practice on the grounds of the fact that the resulting benchmark impulse response function does not necessarily emerge from a single model. In fact, it would only be the case if $R_k^{pq}(j)$ was monotonous in the rotation angle $\theta(j)$ at each k. But since there is no guarantee for monotonicity, the different points in the benchmark impulse response function will generally come from different models, i.e. different values of $\theta(j)$. The issue is the same across all variables and all shocks. As a consequence, the identified shocks are no longer orthogonal to each other and we present a set of impulse responses which are not simultaneously generated by the same model.

In order to remedy these shortcomings, Fry and Pagan (2007) suggest to choose the rotation angle $\theta(j)^*$ as a benchmark for which the impulse responses are as close as possible to the median response across all shocks and variables. In their approach, the impulse responses are thus produced by a single model and a set of orthogonal shocks, while at the same time somewhat preserving the consensus view that the median is a good way of summarizing the results. As a criterion for choosing the rotation angle $\theta(j)^*$, for which the impulse response functions $R_k^{pq}(j)$ are closest to $med(R_k^{pq})$, the authors suggest to use the sum of the squared deviations, normalized by the standard deviation across the response horizon and across all variables and shocks.

$$\min_{\theta(j)} \sum_{p=1}^{n} \sum_{q=1}^{n} \sum_{k=1}^{h} \left(\frac{R_k^{pq}(j) - med(R_k^{pq})}{std(R_k^{pq})} \right)^2 \tag{5.10}$$

where h is the horizon considered in the impulse response function and $std(R_k^{pq})$ is the standard deviation of R_k^{pq} across all accepted draws. While employing the standard approach of reporting the median response in our benchmark case, we use the approach of Fry and Pagan (2007) as a robustness check for our results. Similarly, we compute the variance decomposition as the median of the variance decompositions produced by all accepted draws, but report the variance decomposition resulting from $\theta(j)^*$ as a robustness check.

5.4 Estimation Results

We now turn to the empirical findings. We estimate the Bayesian VAR described in Section 2. Throughout the analysis, we identify monetary shocks using the sign restrictions presented in Table 1. It is important to emphasize that we do not place any restrictions on the capital flow variables (cap_t) in the definition of the vector of endogenous variables y_t. We therefore allow the data to speak for itself in terms of the responses of our variables of interest. In addition, the real exchange rate ($reer_t$), the trade balance (tb_t) and the relative equity returns ($eq_t - eq_t^*$) are left unrestricted in each of the specifications we employ. This is important because we draw inference upon the correlations between the impulse responses of these and the capital flow variables. We first present impulse response functions of the endogenous variables in the model and later their variance decomposition.

5.4.1 The Response and Composition of Capital Flows

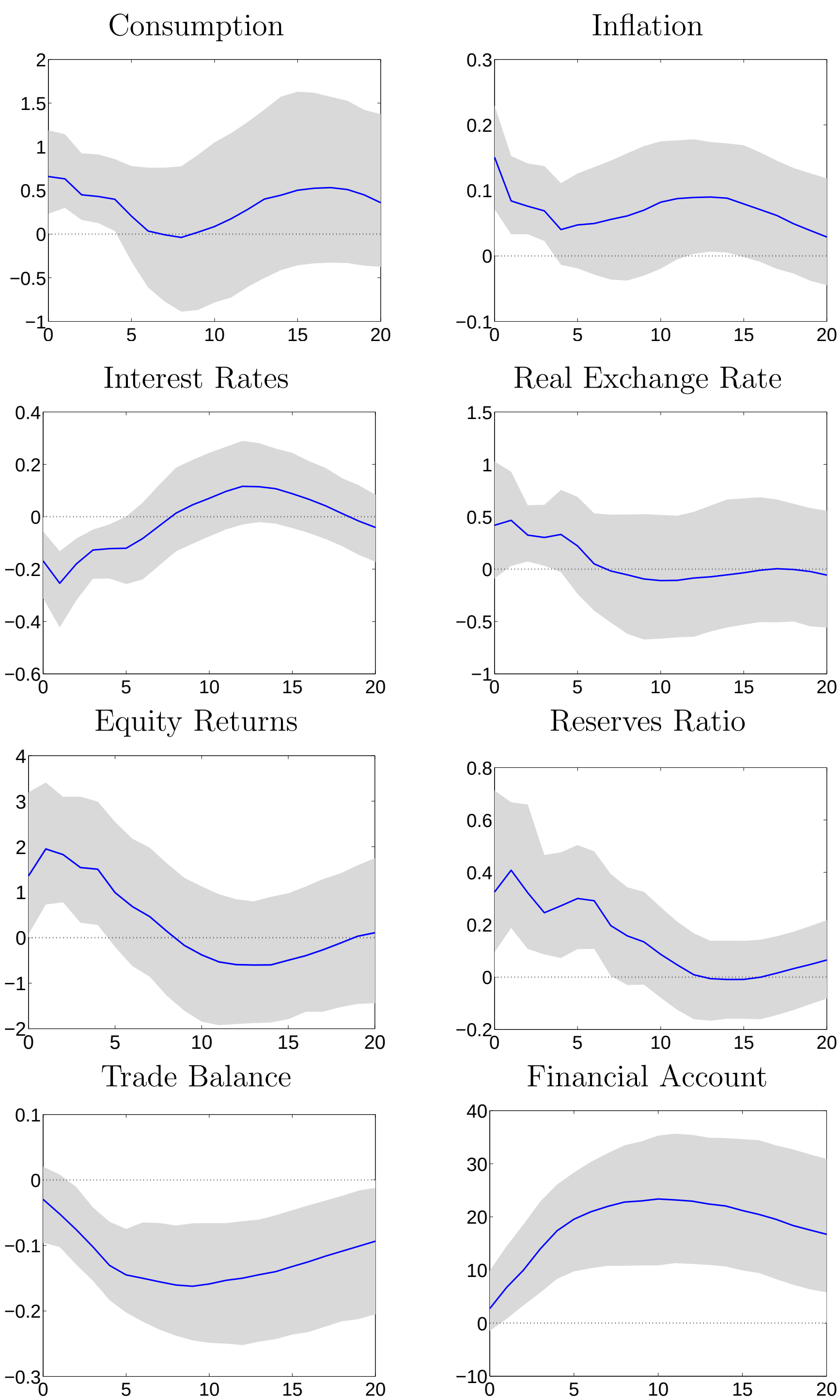
In this subsection, we present the results from estimating the Bayesian VAR presented in equation (5.5). The vector of endogenous variables is defined as in (5.6). The model thus includes seven control variables plus one of the capital flow variables at a time. We estimate the model for each of the four different types of capital flows as well as the financial account as an aggregate. In our benchmark specification, the capital flows are nominal flows denominated in billions of US dollars. Moreover, the capital flow variables are defined such that a positive value signals a net capital inflow into the United States. The fact that these variables can thus take both positive and negative values is the reason why we use the variables in levels instead of log terms. It is clear that the use of nominal flow variables is subject to the critique that eventual impulse responses to the monetary shock can be distorted by the response of relative price levels and/or the exchange rate to the same shock. We will address this caveat in the robustness section.

Figure 5.1 shows the impulse response of all eight variables in the VAR to a one-standard deviation expansionary monetary shock when the financial account, i.e. the net aggregate of all types of capital flows is included in the model. In this benchmark specification, the monetary easing implies a 16 basis points (b.p.) reduction in nominal short term interest rates relative to the rest of the world. The solid line in each of the subplots illustrates the median response of the respective variable to the shock. It is presented for a horizon of 20 quarters. Following most of the sign restrictions literature, we also report 16th and 84th quantiles of the response functions (Uhlig, 2005).[8]

The impulse responses depicted in Figure 5.1 suggest that the reserve ratio nb_t, relative consumption $c_t - c_t^*$ and the percentage difference in CPI inflation $cpi_t - cpi_t^*$ react positively to the monetary shock as imposed via the identification restrictions. However, in particular in the case of $c_t - c_t^*$ the response is short-lived. The interest rate differential $i_t - i_t^*$ falls for few additional periods before its response reverts to zero. The ex ante unrestricted equity return differential $eq_t - eq_t^*$ increases for about ten quarters following the shock. In terms of magnitude, on impact a monetary policy easing shock

[8]In the VAR literature using Cholesky or Blanchard-Quah type decompositions, two standard deviation error bands are typically reported. This is equivalent to using 2.3% and 97.7% quantiles. However, in the sign restrictions literature inference is typically based on a limited number of draws which increases the uncertainty around the quantiles.

Figure 5.1: Benchmark Specification



by 100 b.p. raises relative equity returns by about 6%. The distribution of the impulse responses is strongly positive for about five quarters. Moreover, on impact the one standard deviation expansionary monetary shock leads to a deviation from equity parity of more than 1.5 percent.

After an initial appreciation, the real exchange rate $reer_t$ begins to depreciate in response to the expansionary monetary shock, and then stabilises and appreciates again after about 8 quarters.[9] This pattern is somewhat different, though not inconsistent with the evidence on "delayed overshooting" in the literature, as e.g. shown in Scholl and Uhlig (2008). While its dynamics are in line with the standard UIP reasoning underlying the overshooting model, the magnitudes of the initial appreciation and the rather weak depreciation thereafter are unexpected. An explanation for these findings might be the fact that we use real effective exchange rates (given the purpose of our analysis) rather than bilateral nominal exchange rates as is standard in the work on testing Dornbusch's overshooting hypothesis. Moreover, another reason may have to do with the fact that we restrict the consumption differential to rise on impact of the monetary shock. This is a restriction not employed in Scholl and Uhlig (2008). Intuitively speaking, a rise in consumption will put appreciating pressure on the real exchange rate. A robustness exercise presented in Figure A.8 shows that the real exchange rate indeed depreciates for most of the response horizon if we relax this restriction, while the main results for capital flows remain unchanged. Note however that the initial

[9] The link between monetary poliocy shocks and exchange rates is covered extensively by the literature on testing Dornbusch's overshooting hypothsis - see Eichenbaum and Evans (1995), Kim and Roubini (2001), Faust and Rogers (2002) and Scholl and Uhlig (2008).

appreciation of the exchange rate following a monetary easing, which goes against the standard UIP reasoning, is found also in Scholl and Uhlig (2008) and Grilli and Roubini (1995).

The next subplot of Figure 5.1 shows that the trade balance tb_t worsens significantly and persistently throughout the entire response horizon. In terms of magnitude, a monetary policy easing shock by 100 b.p. worsens the US trade balance by about 1% of US GDP after 8 quarters. This finding is in line with Kim (2001), Bems, Dedola and Smets (2007) and Barnett and Straub (2008) who also find evidence in favour of a significant current account worsening following an expansionary monetary shock. The reason is likely to be an "income absorption effect", i.e. an increase in domestic import demand following the fall in the interest rate differential. We therefore provide evidence against the presence of a significant expenditure switching effect, the importance of which is highly disputed in the open economy macroeconomics literature.[10] Given the balance of payments identity, we would expect the net aggregate of all capital flows to react in the opposite way of the trade balance. And indeed we find that the response of the financial account is almost the exact mirror image of the trade balance response. This strong and persistent inflow of capital into the United States in response to the expansionary monetary shock may be an additional explanation for the above discussed response of the real exchange rate.

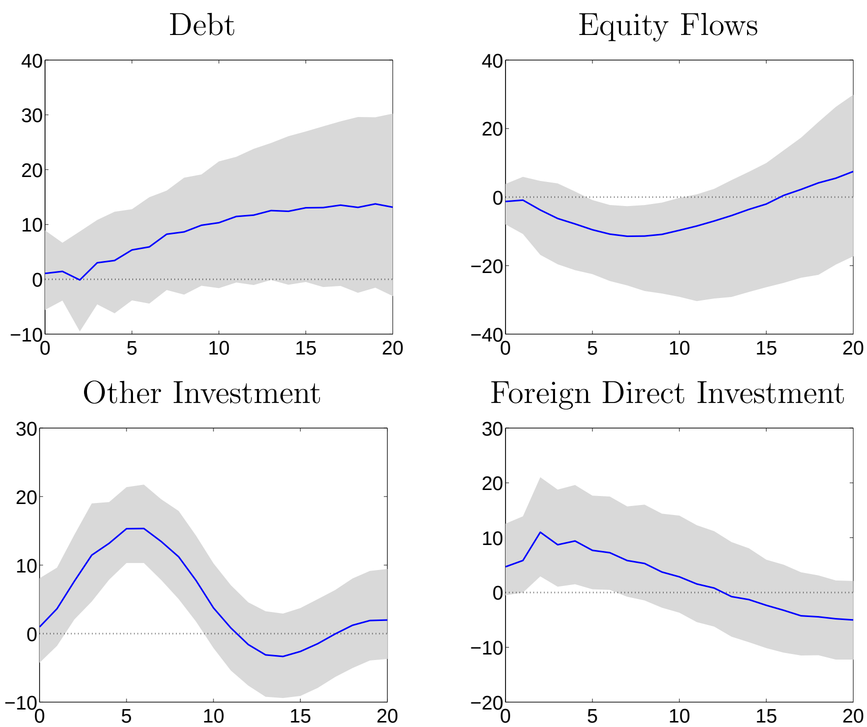
As a next step, we decompose the financial account into its different components in order to better understand the transmission mechanism and the heterogeneity across different types of capital. As discussed above, differ-

[10]For early references see Obstfeld and Rogoff (1995) and Betts and Devereux (1996).

ent types of capital flows have their own characteristics and determinants. While trade credits, loans and currency flows are likely to be at least partly driven by financing motives for trade flows, FDI, equity and debt flows may be driven more by return considerations. We hope to uncover some of these motives by examining the correlation between the impulse responses of our control variables and the responses of each of the different types of capital flows. As outlined in the previous section, the literature on cross-border capital flows has attempted to answer the question of whether return-chasing or portfolio rebalancing motives are dominant in driving cross-border equity flows by establishing correlations between net equity flows and equity returns at particular points in time. The present study has the advantage of allowing for an additional time dimension through the impulse response functions we generate. This allows us to deduce investors' decisions by tracing the response of both capital flows and the relevant macro variables to structural economic shocks along the response horizon.

Figure 5.2 shows the impulse response functions of foreign direct investment (fdi), equity investment flows (*equity*), debt investment flows (*debt*) and other investment flows (*oi*) to an expansionary monetary policy shock. The responses of the control variables are omitted since they do not differ in any noteworthy way from what is shown in Figure 5.1. The solid lines in Figure 5.2 again represent the median impulse response of the respective variable to the monetary shock. As discussed above, Fry and Pagan (2007) challenge the usefulness of the median as a summary measure of the impulse response distributions at each point along the response horizon. In order to show that our results are robust to this criticism, we report the median im-

Figure 5.2: Benchmark Specification: Different Types of Capital Flows



pulse responses together with the impulse responses generated by minimizing the Fry and Pagan (2007) criterion. In Figure A.6, the dashed lines represent the set of response functions produced by the rotation angle θ^*, while the solid lines show the median response. It is obvious that the two sets of response functions look very much alike. In fact, the qualitative results presented below hold irrespectively of the summary measure considered. In order to save space, we therefore concentrate in what follows on the more commonly used median response as a benchmark case.

A first glance at the results of Figure 5.2 reveals that the different types of capital flows react very heterogeneously to monetary policy shocks. The main striking finding is the opposite response of equity flows to that of debt flows, as well as FDI and other investment flows. In order to interpret this finding, it is instructive to recall the response of key control variables in the model. In particular, we witnessed a positive deviation from equity parity for about one to one and a half years following a monetary policy shock. Thus, taking into account the entire response horizon, we confirm the predictions of Hau and Rey (2006) and others (conditional on the monetary shock), namely that a positive deviation from equity parity is associated with portfolio equity outflows. However, the opposite is the case for bonds: the decline in US short-term interest rates - and thus the rise in bond returns - induces net inflows into US bonds, thus implying a positive correlation between bond returns and bond flows that is consistent with a return-chasing motive of investors.

A crucial advantage of our approach is that we can trace the dynamic response pattern of the various variables over time, and in particular we can trace both equity returns and the evolution of net equity flows over time in response. In fact, our results suggest that equity outflows materialise later than the rise in relative equity returns in response to a monetary policy shock. Thus these outflows occur some time after returns have already begun deviating from parity. This suggests a more differentiated answer to the question of whether return chasing or portfolio rebalancing motives dominate investors' decisions. In particular, as regards the return chasing motive, it is important to examine the risk-adjusted performance of an investment strategy of betting on violations of the equity parity condition. In fact, our results suggest that an investor who is overexposed to the US stock market will refrain from rebalancing his portfolio if the expected returns from buying additional US equity are large enough relative to the risk associated with the investment.

In order to illustrate this point, we take the perspective of a Bayesian investor who considers to bet on violations from (conditional) equity parity at each point during the response horizon. Note that the impulse response of the equity return differential k periods after the shock is the period k excess return due to the monetary policy shock of the following investment strategy: in period $k-1$ after the shock, the investor sells one (foreign currency) unit of foreign equity, exchanges the payoff into US dollars and reinvests it into US equity. Following Scholl and Uhlig (2008), we calculate the implied return-to-risk (Sharpe) ratio of this (conditional) investment as $\frac{eq_t - eq_t^*}{sd(eq_t - eq_t^*)}$ for every point in the response horizon. In particular, we define it as the ratio of the posterior mean excess return and the posterior standard deviation of the distribution of impulse response functions in period k. This measure gives us an idea of the reward and the risk a potential investor faces when betting on a deviation from equity parity at different points along the response horizon.[11] It is important to note that the investor bets on violations from equity parity conditional on the monetary shock. In other words, as Scholl and Uhlig (2008) argue, the implied Sharpe Ratio we construct for the hedging strategy presented takes the perspective of a Bayesian investor, who remains uncertain about the precise impact of the monetary shock, but can insure against any other types of shocks that might occur during the investment horizon.

As Figure A.7 illustrates, the Sharpe Ratio increases strongly on impact of the shock and then falls persistently. It reaches a value of about 0.75 at the point at which substantial amount of net equity investment flows out of the US. The reward-to-variability ratio of betting on positive deviations from equity parity thus needs to fall by a substantial amount before investors start rebalancing their portfolio and selling US equity. In sum, the portfolio rebalancing motive does trigger an outflow of equity from the country in which relative equity returns rise. However, the outflow only occurs when the profitability of chasing higher expected returns diminishes to a sufficient extent.

Figure 5.2 also shows the response of foreign direct investment to the monetary easing. Although most of the relevant literature has focused on equity flows when discussing the implications of the portfolio balance model

[11] Scholl and Uhlig (2008) compute Sharpe ratios for a Bayesian investor betting on deviations from uncovered interest rate parity in response to a monetary policy shock. In contrast to the present paper, the authors focus on the return-to-risk ratio of bets of differing length that all begin in the impact period of the shock. For the purpose of this study, however, it is more interesting to examine simple one-period bets that begin at different points in the response horizon.

for cross-border capital flows, similar arguments can be made for FDI. The figure shows that there is a net inflow of FDI immediately following the shock. The inflow is quite persistent and remains substantially positive for about two years. This result is quite intriguing given the fact that equity flows respond in the opposite way. As discussed previously, the inflow of FDI coincides with simultaneous increases in the equity return differential $eq_t - eq_t^*$ and the consumption differential $c_t - c_t^*$. This implies that, contrary to equity flows but similar to bond flows, foreign direct investment appears to be driven by expected returns rather than portfolio rebalancing considerations. A reason for this finding might be the fact that short term risk and portfolio balance considerations play less of a role for FDI flows than for equity flows as the former are typically more long term oriented. In any case, the result emphasizes the particular nature of FDI flows compared to less concentrated forms of equity investment.

The third subplot in Figure 5.2 presents the impulse response of debt flows to the monetary shock. The responses of the interest rate differential, the inflation differential and the real effective exchange rate at least allow for a suggestive interpretation of the response of debt instruments to a deviation from interest rate parity. In particular, it appears that investment in debt increases quite persistently in the US relative to the rest of the world when US interest rates fall, thus inducing a deviation from interest rate parity. Hence, contrary to portfolio equities, the monetary policy shock induces a positive conditional correlation for bonds between returns and flows - as returns rise when interest rates fall - suggesting that return chasing motives play a dominant role for debt flows. A stronger argument in this respect could be made if bilateral nominal exchange rates were included in the model such that deviations from interest rate parity could be observed and a similar analysis as in the case of equity investment could be conducted. However, due to the rest of the world definition behind the construction on the real exchange rate, this is not possible such that the evidence must remain suggestive.

The last plot in Figure 5.2 shows the response of other investment to the monetary easing, showing that other investment flows into the United States increase strongly and significantly during the first part of the response horizon and become insignificant thereafter. In order to understand the reasons behind this finding, it is perhaps useful to remember that major categories of these flows are trade credits, loans and currency flows. These types of capital are typically used to directly finance import expenditure. Hence, one might categorize these flows as borrower rather than investor driven. It is then reasonable to expect that inflows of this type of capital should occur prior to the import expenditure actually being made. And this is precisely what we observe.

In summary, the evidence of this section has shown that monetary policy easing shocks cause *net inflows* in debt securities, foreign direct investment (FDI) and other investment, while inducing *net outflows* in portfolio equities from the United States. Monetary policy shocks thus entail a conditional negative correlation between flows in portfolio equity and debt. A key for understanding this conditional correlation is the effect of monetary policy shocks on asset price returns, which induces a positive correlation between equity returns and bond returns. Overall, our evidence suggests that, conditional on monetary policy shocks, a portfolio rebalancing motive dominates for investment decisions in equity securities but a return chasing motive is the main driver for investments in bonds.

5.4.2 Variance Decomposition

As a complement to the analysis in the previous subsection, we decomposed the variance of the endogenous variables in our model in order to determine the variance share explained by the monetary shock. Notice that the findings presented here are based on the benchmark specification but are not sensitive to identifying additional (aggregate supply and demand) shocks.[12] Table 5.1 contains the median results for the capital flow variables of interest and the trade balance. A first glance at the numbers suggests that monetary policy shocks are important drivers of all these variables. The share of the variation explained by the monetary shock ranges from 4 to 23 percent across horizons of one to four years. Another compelling finding is that the monetary policy shock explains around 20 percent of the variance of both the trade balance and the financial account at various horizons while the explanatory power for the disaggregated capital flows is much more limited in size. This finding may seem contradictory at first but is in line with the fact that the variance of the financial account variable is the sum of both the variances and the covariances of the individual capital flow variables. It appears that the monetary shock explains the covariances very well, while this is less the case for the individual variances of the flows. Intuitively, it is reasonable that the monetary shock must have strong explanatory power for the financial account as an aggregate if it does so for its counterpart in the balance of payments identity.

[12] The resulting variance shares for these additional shocks are available upon request.

Table 5.2: Variance Decomposition

Horizon	Financial Account	FDI	Equity	Debt	Other	TB
1 Year	12.6	7.3	4.1	3.5	14.0	16.7
2 Years	20.7	8.5	7.4	5.7	22.9	22.9
3 Years	22.2	8.7	8.6	7.5	20.5	22.4
4 Years	22.3	9.0	8.3	8.8	19.2	21.5

Table 5.3: Variance Decomposition: Fry and Pagan (2007)

Horizon	Financial Account	FDI	Equity	Debt	Other	TB
1 Year	10.2	17.2	2.6	6.6	9.4	14.6
2 Years	19.4	12.8	2.8	6.2	21.7	17.0
3 Years	20.7	10.6	4.0	8.9	19.8	20.2
4 Years	24.7	9.0	4.1	12.1	18.0	22.4

Most of the individual types of capital flows, on the other hand, are purely incentive driven. Their variances can thus be rather decoupled from the financing needs for trade expenditure as long as their covariances are such that the balance of payments accounting identity is achieved. In other words, the investor does not consider the financing needs of the country he or she invests in. It is only in the aggregate that investors' decisions need to be such that capital flows balance the net flow of goods and services. This argument is perhaps emphasized by the fact that other flows are the only individual type of capital that is driven by monetary shocks to a similar extent as the trade balance. As we outlined above, these types of capital are typically less incentive driven and likely to be very closely aligned with changes in the trade balance.

In Table 5.3, we present the results from an equivalent variance decomposition based on the set of impulse responses that minimize the Fry and Pagan (2007) criterion. The numbers show that the qualitative arguments we made above are not sensitive to a different summary measure of the distribution of impulse response functions. At the same time, however, it is interesting to note that the precise numbers for each individual variable differ quite strongly in some cases. This suggests that caution is in order when interpreting variance decompositions solely based on one summary measure of the distribution of impulse response functions.

5.4.3 Robustness Analysis

We conduct a battery of robustness checks to ensure that the main findings in the previous sections are not sensitive to the specification of the empirical model. In this subsection, we present the results obtained from these tests.

In the previous section, we employed a restriction in the identification scheme of a monetary policy shock, which differs from the analysis of Uhlig(2005). In particular, we assume that an expansionary monetary shock must have a positive effect on consumption. The reasoning behind this assumption is rather obvious and it is well-established in the literature. We believe that it helps to identify monetary shocks with greater precision. However, one might argue that the restriction implies an unnecessary reduction of the degrees of freedom in the empirical model. As a first robustness check, we therefore identify the monetary policy shock solely on the basis of the remaining three restrictions, i.e. the restrictions on the response of the interest rate differential, the inflation differential and the reserve ratio. The resulting responses of the endogenous variables in the model can be found in Figure A.8 for the case in which the financial account is added to the basic specification. The impulse response functions presented show that the consumption differential still reacts positively to the expansionary shock in the impact period. Following a brief initial appreciation, the real exchange rate depreciates strongly in response to the monetary shock and shows evidence of delayed overshooting. The impulse responses of the remaining control variables and the financial account do not change in any important way compared to the benchmark case. Figure A.9 shows that the same is true for the responses of the other capital flow variables.

It has frequently been argued that the number of shocks identified in a VAR is positively related to the probability of identifying each individual shock correctly (Paustian, 2007). As a second robustness check, we therefore identify two additional structural shocks simultaneously with the monetary shock. We have chosen simple aggregate supply and demand shocks because the underlying identifying restriction are rather uncontroversial. In particular, Table 5.1 shows that we require the aggregate supply shock to reduce inflation and to have a positive effect on consumption whereas an aggregate demand shock must increase the interest rates, inflation and consumption. Formally, we extend the method outlined in Section 5.3 by requiring that a candidate draw of the decomposition of the variance covariance matrix must, in order to be accepted, uncover one shock that obeys the restrictions of the monetary shock, one that obeys the restrictions of the aggregate supply shock and one that obeys the restrictions of the aggregate demand shock. The resulting impulse response functions for the capital flow variables are shown in Figure A.10. It is immediately obvious that the response functions do not differ in any important way from their equivalents in the benchmark specification.[13] If anything, the responses are more precise than before.

In this study we are interested in the channels through which the adjustment of the financial account takes place following the occurrence of a structural shock. The comovement of capital flows with the equity return differential $eq_t - eq_t^*$ naturally plays an important role in this context. We have so far used the equity return differential in US dollars in our model. The reason is that changes in relative equity returns should only play a role for the re-allocation of capital across borders is they are not offset by exchange rate movements. However, in order to ensure that the reaction of $eq_t - eq_t^*$ to the expansionary monetary shock is not entirely due to exchange rate fluctuations and indeed reflects asymmetric equity price changes, it is instructive to include the equity differential in local currencies as a robustness check. As Figures A.11 and A.12 show, the impulse response functions for the different types of capital flows and the equity return differential look qualitatively very similar to the benchmark results and deliver the same set of qualitative conclusions. The only striking difference is that the equity return differential increases for about eight instead of five quarters after the shock impacts the economy. In line with the above reasoning, this result perhaps strengthens the view that portfolio rebalancing motives are an important driving factor in the cross-border allocation of equity investment.

In the above analysis, we have used unadjusted capital flow variables. The

[13]In order to save space, we have not reported the responses of the control variables themselves. But their impulse responses do not change either. The results are available upon request.

results are therefore subject to the criticism that the response of the variables to the monetary shock might be driven by movements in relative price levels or the exchange rate. We tackle this criticism by including our capital flow variables as ratios to GDP as a robustness check. The resulting impulse response functions for each of the capital flow variables are contained in Figure A.13. We can see that, compared to the benchmark specification, the adjustment by GDP lowers the impulse responses of all variables during the first quarters after the shock. The reason is simply that consumption and GDP increase in response to the shock for a few periods. Hence, positive responses become weaker whereas negative responses become more pronounced. However, the fact that all of the response functions retain their qualitative shape suggests that the impulse responses of the unadjusted capital flow variables in our benchmark specification are not simply the result of changes in relative price levels or the exchange rate.

Up until now, we have included one capital flow variable in the model at a time. However, it is clear that there might be important interdependency between the different types of capital flows that we miss if the remaining variables are omitted. As a robustness check, we therefore augment the VAR by dimension three such that the four different types of capital flows can be included at the same time. In spite of the fact that we are now working with a VAR of dimension eleven, the resulting impulse response plots shown in Figure A.14 look almost exactly the same as in our benchmark specification.[14] There appears to be no loss in the precision of the estimates, which might be due to the fact that the interlinkages between the capital flow variables are

[14]The response of the reserve ratio is omitted for presentation purposes.

indeed important and improve the fit of the model.

We also checked for the robustness of our results to an alternative definition of the "Rest of the World". We now define the "Rest of the World" as the G7 countries plus a range of additional economies, the selection of which was made solely subject to data availability.[15] Figure A.15 shows that the results are robust to this redefinition.

Finally, we considered one potential criticism to our results with regard to the dynamics of net debt flows. In particular, the interest rate we have considered so far has been the short-term money market rate, while the return on international debt flows is rather better represented by the evolution of long-term interest rates. Therefore, in this exercise we replace the differential of short-term money market interest rate by the 10-year bond yield differential between the United States and other G7 economies. As shown in Figure A.16, our main results that (i) debt flows and debt returns are positively correlated and (ii) debt and equity flows are negatively correlated still hold, confirming the dominance of the *return chasing* motive behind international debt flows following a monetary policy shock.

5.5 Discussion

The evidence of the paper has shown that monetary policy shocks exert a substantial effect on the dynamics and composition of US capital flows. In the aggregate, a monetary policy easing shock of 100 basis points leads to net capital inflows and a trade balance deficit of about 1% of US GDP after 8

[15] The "Rest of the World" now includes the G7 (minus the US) as well as Australia, Belgium, Denmark, Finland, Netherlands, Norway, Spain Sweden, and Switzerland.

quarters. The key finding of the paper is that monetary policy shocks induce a negative conditional correlation between flows into equities and bonds, while causing a positive conditional correlation between equity returns and bond returns.

Moreover, for equities there is a negative conditional correlation between flows and returns, i.e. a rise in equity returns in response to monetary policy shocks is eventually associated with an outflow in equity portfolio investment from the country. The opposite is the case for bonds, for which there is a positive conditional correlation between returns and flows.

Yet it is not only the direction of capital flows and returns that exhibit an intriguing pattern, but also the dynamics of flows and returns. While returns - interest rate differentials and relative equity returns - react instantaneously to monetary policy shocks, capital flows react much more gradually over time, with their peak response occurring only after about eight quarters or more. The strength of the methodology of the paper is hence that it allows tracing not only the overall reaction of capital flows and returns, but also to understand how the dynamics of these responses differ.

A central objective of the paper has been to contribute to the literature on the determinants of portfolio choice, and how asset price movements are related to such portfolio decisions. The evidence of the paper is consistent with a portfolio rebalancing motive for equity securities, and a motive akin to return chasing for bonds.

An intriguing issue is that flows and returns in equities and in bonds respond in the opposite way to such monetary policy shocks. The literature on asset price comovements has found it hard to explain the asset price comovements empirically, in particular the strong time variations in stock-bond return correlations. The findings of the paper suggest that such a positive correlation between stock returns and bond returns is present precisely when discount rate effects dominate. Of course, it also implies that the correlation may be very different when other shocks dominate. The present paper thus focuses on one specific shock for portfolio choice and asset prices, while we leave it for future research to condition the analysis on other types of economic shocks.

Chapter 6

Conclusions

A growing number of both theoretical and applied economists have devoted their time to the study of the conduct of monetary policy in recent years. This thesis has contributed to different fields of this literature.

Chapter 2 is motivated by the idea that inflation targeting differs from other disinflation policies in important respects. In particular, a strict interpretation of an inflation target allows the policymaker to tolerate only minor deviations from target. But adjusting the policy instrument such that the inflation rate is reduced to a new target and defending this target rigorously must preserve inflationary distortions such as wage and price differentials to an exceptional degree. The chapter employed a dynamic general equilibrium with Taylor wage contracts to show that the use of strict inflation targeting as a disinflation policy may result in a slump in output and a considerable increase in macroeconomic volatility. Important determinants of the magnitude of the macroeconomic oscillations in the post-disinflation state turned out to be the size of the reduction in the inflation rate and the degree of returns to labour. In Chapter 3, the analysis was extended to an open-economy setting demonstrating that the exchange rate can act as a stabilizer by effectively relieving wages of part of the burden of reducing the inflation rate. The more the economy is open, the smaller the magnitude of the macroeconomic oscillations will be after the disinflation policy is applied.

Chapter 4 of this thesis employed a Markov switching framework to allow for an alternative characterization of macroeconomic news effects on the foreign exchange market. The choice of the model was motivated by the presumption that monetary policy announcements do not simply affect the foreign exchange market as shocks to an otherwise continuous process. On the contrary, news effects may change the entire data generating process underlying a market's dynamics. An econometric specification allowing for regime switches therefore appears appropriate. Indeed, one particular benefit of applying such a model is that it facilitates a plausible interpretation of observed nonlinearities. We found strong evidence for nonlinear regime switching between a high-volatility "informed trading" state and a low-volatility "liquidity trading" state driven by monetary policy announcements that come as a surprise to the market. We also uncovered significant market positioning prior to the announcement, indicating that market participants limit their risk exposure just before policy decisions are made.

Chapter 5 investigated the impact of monetary shocks on the direction and the composition of international capital flows. The chapter employed a standard structural VAR specification to identify monetary policy shocks by means of the pure sign restrictions approach used in Canova and De Nicolo (2002) and Uhlig (2005). The empirical analysis yielded two key findings. First, US monetary policy shocks exert a statistically and economically meaningful effect on US capital flows and the trade balance. An exogenous easing of US monetary policy by 100 basis points induces net capital inflows and a worsening of the US trade balance of around 1% of GDP after 8 quarters. The second main result focuses on the effect of monetary policy shocks on the composition of US capital flows. Intriguingly, it is found that an exogenous US monetary policy easing causes net inflows in debt securities, foreign direct investment (FDI) and other investment, while inducing net outflows in portfolio equities from the United States. Monetary policy shocks thus entail a conditional negative correlation between flows in portfolio equity and debt. A key for understanding this conditional correlation is the effect of monetary policy shocks on asset prices. While a monetary policy easing implies a decrease in short-term (and long-term) interest rates, it also causes the above mentioned increase in relative equity returns. Overall, our evidence suggests that monetary policy shocks induce negative conditional correlations between flows in bonds and equity securities. Moreover, they cause a negative conditional correlation between equity flows and equity returns, and and a positive conditional correlation between bond flows and bond returns.

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APPENDIX

A.1 Appendix for Chapter 2

A.1.1 Log-linear Definitions and Equations

In the following equations lower-case symbols denote log deviations of variables from their reference steady state values, i.e. $v_t = log\frac{V_t}{V_R}$:

$$\mu_{t+1} = m_{t+1} - m_t \tag{A.1}$$

$$s_t = p_t + c_t \tag{A.2}$$

$$m_t - p_t = c_t - \frac{\beta}{1-\beta} i_t \tag{A.3}$$

$$s_{t+1} - s_t = i_t \tag{A.4}$$

$$y_t = \frac{\sigma}{1-\sigma}(p_t - w_t) \tag{A.5}$$

$$\begin{aligned} x_t = \frac{1}{1+\varepsilon(\zeta-1)}[&\frac{1}{1+\beta}[s_t + \varepsilon(\zeta-1)w_t + (\zeta-1)n_t] \\ &+\frac{\beta}{1+\beta}[s_{t+1} + \varepsilon(\zeta-1)w_{t+1} + (\zeta-1)n_{t+1}]] \end{aligned} \tag{A.6}$$

$$w_t = \frac{1}{2}x_t + \frac{1}{2}x_{t-1} \tag{A.7}$$

$$n_t = \frac{1}{\sigma}y_t \tag{A.8}$$

A.1.2 The Perfect Foresight Solution

In the following, I show in a first step that there is a unique perfect foresight solution for the law of motion of the economy in the post-disinflation state when either $\sigma < 1$ or $\alpha < 1$. In a second step, I show that the sign of the smaller eigenvalue depends on the degree of returns to labor σ and the degree of openness of the economy $1 - \alpha$ as well as the parameters ζ and ε. In the special case of $\alpha = 1$ and $\sigma = 1$ a saddlepath equilibrium does not exist. The proofs are outlined for the open-economy setting but hold equivalently in the closed economy setting if α is set to unity.

Stability

As long as $\alpha < 1$ or $\sigma < 1$, the system of equations determining the law of motion of the economy in the post-disinflation steady state has exactly one stable eigenvalue such that there is a unique perfect foresight solution. To see this, notice that the law of motion of the economy is given by

$$\phi_{t+1} - \frac{1+\beta}{\beta}\frac{1-\alpha\sigma+\gamma}{1-\alpha\sigma-\gamma}\phi_t + \frac{1}{\beta}\phi_{t-1} = \mu_D(\frac{1-\beta}{\beta} - \frac{2\gamma}{1-\alpha\sigma-\gamma}) \tag{A.9}$$

The characteristic equation of this 2nd order difference equation is given by

$$\omega^2 + b\omega + c = 0 \tag{A.10}$$

where $b = -\frac{1+\beta}{\beta}\frac{1-\alpha\sigma+\gamma}{1-\alpha\sigma-\gamma}$ and $c = \frac{1}{\beta}$. Notice that in the special case of $\alpha = 1$ and $\sigma = 1$, the two eigenvalues of this difference equation are readily derived as $\lambda_1 = -1$ and $\lambda_2 = -\frac{1}{\beta}$, implying that one eigenvalue lies on the unit circle and the economy will never converge. In order to prove stability for all other cases, I apply standard results from the continuous time case. In particular, I define the variable z such that $z = \frac{\omega-1}{\omega+1}$ and $\omega = \frac{1+z}{1-z}$.[1] I therefore transform the equation to

$$(1 - b + c)z^2 + 2(1 - c)z + (1 + b + c) = 0 \tag{A.11}$$

Applying standard results for a continuous time characteristic equation, the following condition must hold for the characteristic equation to have exactly one stable eigenvalue.

$$\frac{1 + b + c}{1 - b + c} < 0 \tag{A.12}$$

After some manipulations we have

[1] A stable eigenvalue in the continuous time case has a negative real part. It can be shown that ω lies in the unit circle if and only if z has a negative real part.

$$\frac{1+b+c}{1-b+c} = -\frac{\gamma}{1-\alpha\sigma} \tag{A.13}$$

And since $\alpha, \sigma < 1$, it is clear that this expression is strictly negative. We can conclude that for any parameterization within the restrictions given, there will be exactly one stable eigenvalue. This implies that there is a unique perfect foresight solution to the model and the economy ultimately converges to a new steady state. In the special case of $\alpha = 1$ and $\sigma = 1$ a saddlepath equilibrium does not exist.

The Sign of the Smaller Eigenvalue

I call $p(\omega)$ the 2nd order characteristic polynomial function in ω, i.e. the LHS of the characteristic equation. I have shown that there is one and only one stable eigenvalue when $\alpha < 1$ or $\sigma < 1$. This implies that the characteristic function cuts the horizontal axis precisely once within the range $(-1, 1)$. Now, suppose that the stable eigenvalue is of a positive sign. This implies that $p(0)$ and $p(1)$ must be of opposite sign. Therefore, for the unique stable eigenvalue to be of a positive sign, the following condition must hold:

$$\frac{c}{1+b+c} < 0 \tag{A.14}$$

After some manipulations

$$1 - \frac{1-\alpha\sigma+\gamma}{1-\alpha\sigma-\gamma} < 0 \tag{A.15}$$

Noticing that multiplying by the denominator of the fraction might involve multiplying by a negative number, I find that the condition is fulfilled if and only if $1 - \alpha\sigma - \gamma > 0$, where $\gamma = \frac{\zeta}{1+\varepsilon(\zeta-1)}$. This implies that the unique and stable eigenvalue will be negative for sufficiently high α and σ and sufficiently small ζ and ε.

A.2 Appendix for Chapter 3

A.2.1 Log-linear Definitions and Equations

In the following equations lower-case symbols denote log deviations of variables from their reference steady state values, i.e. $v_t = log\frac{V_t}{V_R}$:

$$\mu_{t+1} = m_{t+1} - mu \tag{A.16}$$

$$s_t = p_t + c_t \tag{A.17}$$

$$m_t - p_t = c_t - \frac{\beta}{1-\beta} i_t \tag{A.18}$$

$$s_{t+1} - s_t = i_t \tag{A.19}$$

$$e_{t+1} - e_t = i_t \tag{A.20}$$

$$y_{Nt} = \frac{\sigma}{1-\sigma}(p_{Nt} - w_t) \tag{A.21}$$

$$y_{Tt} = 0 \tag{A.22}$$

$$c_{Nt} = s_t - p_{Nt} \tag{A.23}$$

$$c_{Tt} = s_t - e_t \tag{A.24}$$

$$p_t = \alpha p_{Nt} + (1-\alpha)e_t \tag{A.25}$$

$$\frac{1}{\alpha}y_t = y_{Nt} \tag{A.26}$$

$$\tau_t = -c_{Tt} \tag{A.27}$$

$$\begin{aligned} x_t = \frac{1}{1+\varepsilon(\zeta-1)}[&\frac{1}{1+\beta}[s_t + \varepsilon(\zeta-1)w_t + (\zeta-1)n_t] \\ &+\frac{\beta}{1+\beta}[s_{t+1} + \varepsilon(\zeta-1)w_{t+1} + (\zeta-1)n_{t+1}]] \end{aligned} \tag{A.28}$$

$$w_t = \frac{1}{2}x_t + \frac{1}{2}x_{t-1} \tag{A.29}$$

$$n_t = \frac{1}{\sigma}y_{Nt} \tag{A.30}$$

A.2.2 Impossibility of a Partial Disinflation

I show in a first step that $\frac{X_0}{X_{-1}} = \frac{X_2}{X_1}$ holds only in the special case of $\mu_D = 1$. This implies that in all other cases the economy is not in the post-disinflation state characterized by perpetual oscillations immediately after the policy is applied. The second step is to show that there is no convergence process of the ratio $\frac{X_t}{X_{t-1}} = \frac{X_{t+2}}{X_{t+1}}$ to its steady state $\frac{X_t}{X_{t-1}} = \mu_D$ thereafter if $1 < \mu_D < \mu_I$. The steady state is unstable and the disinflation policy is infeasible.

Proposition A.2.1. *The equality* $\frac{X_0}{X_{-1}} = \frac{X_2}{X_1}$ *holds in the post-disinflation state only in the special case of* $\mu_D = 1$.

Proof. The ratio of the wage index in period t and its first lag is given by

$$\frac{W_t}{W_{t-1}} = \frac{[\frac{1}{2}X_t^{1-\varepsilon} + \frac{1}{2}X_{t-1}^{1-\varepsilon}]^{\frac{1}{1-\varepsilon}}}{[\frac{1}{2}X_{t-1}^{1-\varepsilon} + \frac{1}{2}X_{t-2}^{1-\varepsilon}]^{\frac{1}{1-\varepsilon}}} \tag{A.31}$$

Setting $\frac{W_t}{W_{t-1}} = \mu_D$ and manipulating yields

$$\frac{X_t}{X_{t-1}} = (\mu_D^{1-\varepsilon}(1 + (\frac{X_{t-2}}{X_{t-1}})^{1-\varepsilon}) - 1)^{\frac{1}{1-\varepsilon}} \tag{A.32}$$

This implies that we have

$$\frac{X_0}{X_{-1}} = (\mu_D^{1-\varepsilon}(1 + \mu_I^{\varepsilon-1}) - 1)^{\frac{1}{1-\varepsilon}} \tag{A.33}$$

$$\frac{X_1}{X_0} = (\mu_D^{1-\varepsilon}(1 + (\frac{X_{-1}}{X_0})^{1-\varepsilon}) - 1)^{\frac{1}{1-\varepsilon}} \tag{A.34}$$

$$\frac{X_2}{X_1} = (\mu_D^{1-\varepsilon}(1 + (\frac{X_0}{X_1})^{1-\varepsilon}) - 1)^{\frac{1}{1-\varepsilon}} \tag{A.35}$$

etc. I will now show that for a given ε the equality $\frac{X_0}{X_{-1}} = \frac{X_2}{X_1}$ only holds if $\mu_D = 1$ and/or if $\mu_D = \mu_I$.

Plugging (A.34) into (A.35), we have

$$\frac{X_2}{X_1} = (\mu_D^{1-\varepsilon}(1 + (\mu_D^{1-\varepsilon}(1 + (\frac{X_{-1}}{X_0})^{1-\varepsilon}) - 1)^{-1}) - 1)^{\frac{1}{1-\varepsilon}} \tag{A.36}$$

Now using (A.33),

$$\frac{X_2}{X_1} = (\mu_D^{1-\varepsilon}(1 + (\mu_D^{1-\varepsilon}(1 + (\mu_D^{1-\varepsilon}(1 + \mu_I^{\varepsilon-1}) - 1)^{-1}) - 1)^{-1}) - 1)^{\frac{1}{1-\varepsilon}} \tag{A.37}$$

Now I set the RHS of (A.37) equal to the RHS of (A.33). I obtain

$$\mu_D^{1-\varepsilon}(1 + \mu_I^{\varepsilon-1}) = \mu_D^{1-\varepsilon}(1 + (\mu_D^{1-\varepsilon}(1 + (\mu_D^{1-\varepsilon}(1 + \mu_I^{\varepsilon-1}) - 1)^{-1}) - 1)^{-1}) \tag{A.38}$$

and finally after some simplification

$$\mu_I^{1-\varepsilon} + 1 = \mu_D^{1-\varepsilon} + (1 + \mu_I^{\varepsilon-1} - \mu_D^{\varepsilon-1})^{-1} \tag{A.39}$$

For a given ε, there are only two positive real valued solutions to this equation. These are $\mu_D = 1$ and $\mu_D = \mu_I$. This proves the above proposition.

□

Proposition A.2.2. *If* $1 < \mu_D < \mu_I$*, then the steady state* $\frac{X_t}{X_{t-1}} = \mu_D$ *is locally unstable and it is impossible to carry out the disinflation policy. If* $0 < \mu_D < 1$*, then the steady state* $\frac{X_t}{X_{t-1}} = \mu_D$ *is locally stable and the disinflation policy can be carried out.*

Proof. From equation (A.32) in Proposition A.2.1 we have that

$$(\frac{X_t}{X_{t-1}})^{1-\varepsilon} = \mu_D^{1-\varepsilon}(1 + (\frac{X_{t-1}}{X_{t-2}})^{\varepsilon-1}) - 1 \tag{A.40}$$

I define $(\frac{X_t}{X_{t-1}})^{1-\varepsilon} = r_t$. Then we have the first order non-linear difference equation

$$r_t = \mu_D^{1-\varepsilon} - 1 + \frac{\mu_D^{1-\varepsilon}}{r_{t-1}} \tag{A.41}$$

This difference equation has a steady state at $r = \mu_D^{1-\varepsilon}$ or $r = -1$. A negative value of r is economically meaningless. Hence, the only economically meaningful steady state is $r = \mu_D^{1-\varepsilon}$, which implies that $\frac{X_t}{X_{t-1}} = \mu_D$. Furthermore, notice that differentiating r_t with respect to its one period lag and evaluating the derivative at the steady state yields a value of $\frac{1}{\mu_D^{1-\varepsilon}}$, the slope of the phase line at the steady state.

Case 1: $1 < \mu_D < \mu_I$ This case corresponds to a partial disinflation policy. It is illustrated in figure A.1. I begin by noticing that there is a natural initial

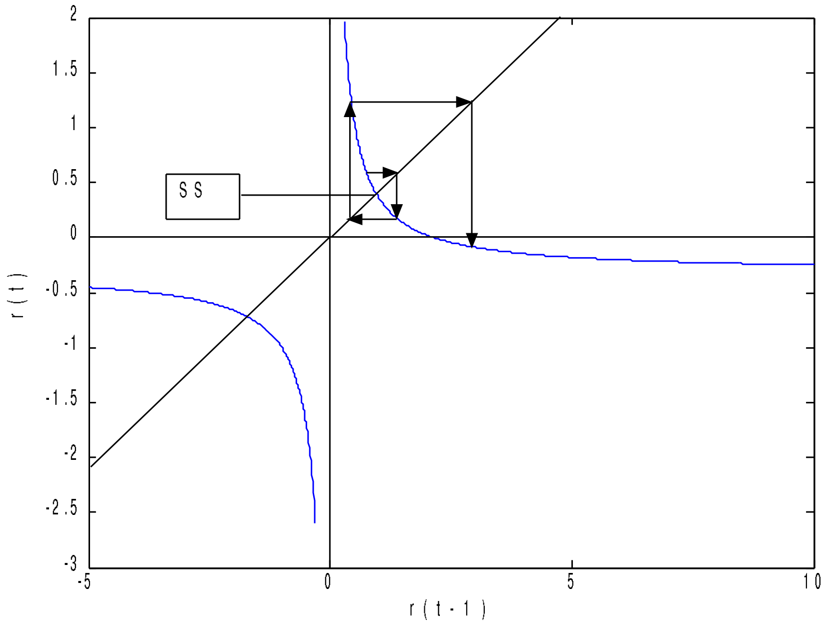


Figure A.1: Case 1: Unstable Pattern

condition for r_t, namely r_0, which is predetermined in this setting. Given this initial value for r_t, the time path of r_t is divergent as the steady state is not locally stable. The reason is that the slope of the phase line given by equation (A.41) is smaller than -1 at the SS. Its horizontal asymptote occurs at a negative value for r_t. This means that eventually $r_t < 0$ which implies that the disinflation policy cannot be carried out. However, notice that our discussion here only covers local as opposed to global stability. As the economy moves farther away from the steady state, and hence nonlinearities become more important, it is possible that the economy behaves in an unexpected way. It is for instance possible that it converges on a so-called 2-cycle, i.e. a path on which fluctuations continue forever at the same amplitude. An analysis of global stability, however, is not within the scope of this paper.

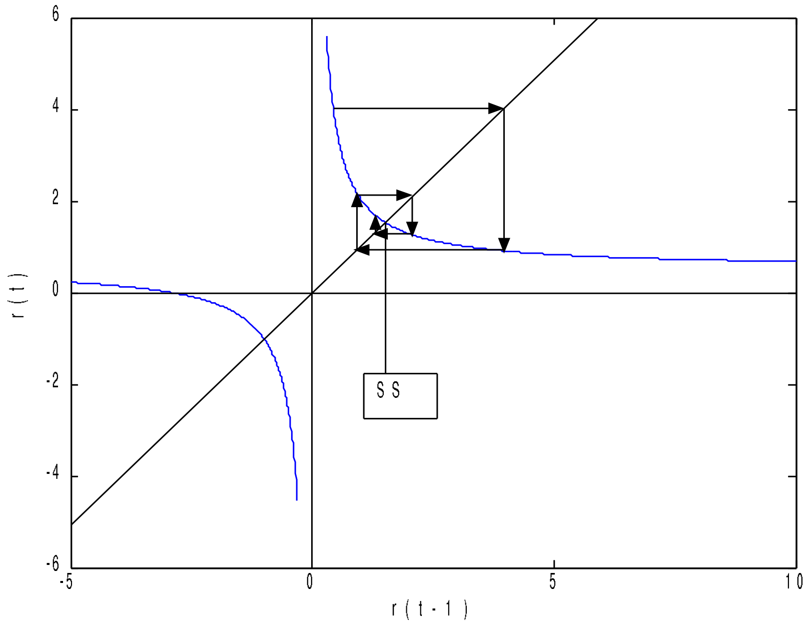


Figure A.2: Case 2: Stable Pattern

Case 2: $0 < \mu_D < 1$ This case corresponds to disinflation policy to a negative inflation rate. It is illustrated in figure A.2. Notice that r_t is locally stable. The reason is that the slope of the phase line given by equation (A.41) is greater than -1 at the SS. Its horizontal asymptote occurs at a positive value for r_t. In contrast to Case 1, this implies that the time path of r_t will follow a stable pattern. The disinflation policy can thus be carried out.[2].

□

Intuitively, the driving factor behind these results is the non(log)linearity of the wage index W_t in X_t and X_{t-1}.[3] This non(log)linearity is due to the fact that labor types are not perfectly complementary but at least partially

[2]As discussed above, notice that I only consider local stability.

[3]If the production function were of Cobb-Douglas type, the wage index would be log-linear and labor types would be less easily substituted than in the present case.

substitutable in the production function. Simply put, if households in one sector set a lower wage than households in the other, the resulting wage index is not given by the simple average of the two wages - as in the linearized version of the model - but by some value smaller than it. The reason is that firms will substitute labor types from the low wage sector for labor types from the high wage sector. Furthermore, firms will find it optimal to engage more strongly into labor substitution the greater is the wage gap between the two sectors and the greater is the elasticity of technical substitution between labor types, ε. This labor substituting behavior of firms is the reason, why it is not possible to keep wage growth at target throughout future periods after the disinflation to a positive rate of inflation. In the following, I will contrast wage setting after (a) a complete disinflation, (b) a disinflation to a positive rate of wage inflation and (c) a disinflation towards a negative rate of wage inflation.

$\mu_D = 1$ is the case of a complete disinflation. I have shown in the main text that a complete disinflation implies that wages set in the post-disinflation state are constant through time for each sector of households. Furthermore, these wages are equal to the respective wage of each sector one period before disinflation. Therefore, this is the one case in which the policy does not introduce dynamics into the behavior of sectoral wages.

Let us now consider the case when $1 < \mu_D < \mu_I$. Before starting the discussion, it is helpful to take a quick look at Figure 2.1 in the main text which illustrates the post-disinflation path of the new wage in the framework of the linearized version of the model. The question is then, why we do not observe equivalent patterns when investigating disinflations to positive new inflation targets in the framework of the nonlinear model.

Let us assume that we are investigating a disinflation that implies a strong enough reduction in the rate of wage inflation for the new wage in period $t = 0$ to be forced to fall below the prevailing new wage that was set in period $t = -1$.[4] In the framework of the linearized model, we would be looking at case 1a in which the new wage set in period zero is lower than in the previous period.

I have chosen the case in which the new inflation target is attained if X_0 takes a value somewhere between X_{-1} and X_{-2}. Now, for a given X_{-1} and X_{-2}, let X_{0_A} be the new wage that would attain the new inflation target in period $t = 0$ in the absence of labor substitution. Notice that $X_{0_A} > X_{-2}$ implies that sectoral wages are less far apart in period $t = 0$ than they were in $t = -1$. Consequently, if we do allow for labor substitution in period $t = 0$, firms substitute less labor away from the high wage sector in period $t = 0$ than they did in period $t = -1$. This implies that the higher wage gains weight in the wage index in period $t = 0$ relative to period $t = -1$ and thereby exerts additional upward pressure on it. This 'labor substitution effect' implies that X_0 must lie below X_{0_A}, the value it would have taken in the absence of labor substitution. Therefore, in terms of log deviations from the reference steady state, the new wage in period zero must be low relative to its counterpart in the framework of the linearized model.

Moving on to period $t = 1$, the smaller is X_0, the greater must agents choose X_1 in order to keep wage growth constant. Furthermore, the higher is X_1, i.e. the greater is the wage gap, the more do firms engage into labor

[4]The reasoning is similar for the opposite case.

substitution away from the higher wage and the higher does X_1 have to be chosen in order to attain the new inflation target. This implies that the 'labor substitution effect' that put downward pressure on X_0 now exerts upward pressure on X_1. And it exerts more upward pressure on X_1, the more downward pressure it exerted on X_0. The fluctuations in the new wage are therefore self-enforcing with the result that the ratio $\frac{X_t}{X_{t-1}}$ falls over time throughout even periods and increases over time throughout odd periods until the new wage hits its zero lower bound in even periods.

$0 < \mu_D < 1$ is the rather unrealistic case of a reduction in wage inflation towards a negative inflation rate. Here, the economy is saddlepath stable. The ratio $\frac{X_t}{X_{t-1}}$ converges towards its post-disinflation steady state $\frac{X_t}{X_{t-1}} = \mu_D$. Notice that this result follows from the above argument as the magnitude of the reduction in the inflation rate is so great that firms engage more strongly into labor substitution in period $t = 0$ than they did in period $t = -1$.

A.2.3 Welfare Analysis

In this section, I analytically derive the welfare implications of the disinflation policy in the special case of a closed economy and constant returns to scale. The reason why I can do so is that closed form solutions of key variables' responses to the application of the policy can be obtained.

I derive the individual lifetime utility for households in both sectors in the CISS, in the POS and in the ZISS without wage dispersion. Social welfare is then some function of the lifetime utility of households in sector A and households in sector B. I begin by deriving agent j's lifetime utility in the initial CISS.[5] We know that aggregate real variables are constant through time in the CISS, while nominal variables grow at the constant rate μ_I. Moreover, consumption and real money demand are equal between households within a sector and, by assumption, also across sectors. It follows that the disutility from work effort ηL_{jt}^{ζ} is the only component of the utility function that is not constant through time and across sectors. In fact, work loads of households in sector A are low in all even periods and high in all odd periods. For sector B, the reverse holds. Making the appropriate substitutions in (2.4), I find that the present discounted value of future utility outcomes of household j, i.e. its lifetime utility, amounts to

$$U_j = \frac{ln(Y_N^{\alpha}) + (1-\delta)ln(\frac{1-\delta}{\delta(1-\beta/\mu_I)}) - \frac{1}{2}\eta(L_{high} + L_{low})}{1-\beta} - \frac{\frac{1}{2}\eta(L_{low}^{\zeta} - L_{high}^{\zeta})}{1+\beta} \tag{A.42}$$

for households in sector A and to

$$U_j = \frac{ln(Y_N^{\alpha}) + (1-\delta)ln(\frac{1-\delta}{\delta(1-\beta/\mu_I)}) - \frac{1}{2}\eta(L_{high} + L_{low})}{1-\beta} - \frac{\frac{1}{2}\eta(L_{high}^{\zeta} - L_{low}^{\zeta})}{1+\beta} \tag{A.43}$$

for households in sector B, where L_{high} is the labor supplied by the high wage sector and L_{low} the labor effort of the low wage sector in any given period and where

$$L_{high} = Y_N^{1/\sigma}(\tfrac{1}{2} + \tfrac{1}{2}\mu_I^{1-\varepsilon})^{\frac{\varepsilon}{1-\varepsilon}}$$
$$L_{low} = Y_N^{1/\sigma}(\tfrac{1}{2} + \tfrac{1}{2}\mu_I^{\varepsilon-1})^{\frac{\varepsilon}{1-\varepsilon}}$$

[5] The ZISS without wage dispersion is a special case of the CISS in which $\mu_I = 1$

with Y_N given by (3.11). Notice that households in the two sectors alternate in setting the higher wage. A relatively high wage guarantees them a low work effort in that same period. As I have assumed an insurance scheme that ensures equal consumption and money holdings across sectors, a household is better off during periods in which its wage is relatively high and its labor effort is relatively low. As households in sector A are the ones with the higher wage in the initial time period, their lifetime utility must be marginally greater than that of households in sector B.[6]

I now move on to derive the individual lifetime utility for household j in the post-disinflation POS. Here, wages in both sectors are constant through time where sector A continuously sets a higher wage than sector B such that the work load of households in sector B is continuously higher than in sector A. As consumption and money demand are equal across sectors in every period of time, this implies that individual lifetime utility must be greater for households in sector B than for households in sector A. The particular feature of the POS is that real as well as nominal variables are constant throughout even and throughout odd periods but differ in between odd and even periods. The lifetime utility of household j in the POS amounts to

$$U_j = \frac{ln(Y^{\alpha}_{N_{even}} Y^{\alpha}_{N_{odd}}) + ln(\frac{1-\delta}{\delta}\frac{1}{1-\beta F^{-1}}) + ln(\frac{1-\delta}{\delta}\frac{1}{1-\beta F}) - \eta(L^{\zeta}_{H_{even}} + L^{\zeta}_{H_{odd}})}{2(1-\beta)} + \frac{ln(\frac{Y^{\alpha}_{N_{even}}}{Y^{\alpha}_{N_{odd}}}) + (1-\delta)ln(\frac{1-\beta F}{1-\beta F^{-1}}) - \eta(L^{\zeta}_{H_{even}} - L^{\zeta}_{H_{odd}})}{2(1+\beta)} \quad (A.44)$$

[6]Notice that I am considering a CISS that lasts an infinite number of periods. The initial time period here simply represents the starting point of the welfare calculations.

for households in sector A, where $F = (\frac{1-\beta B^{1+\varepsilon(\zeta-1)}}{B^{1+\varepsilon(\zeta-1)}-\beta})^{\frac{1}{\zeta}}$ and

$$L_{H_{even}} = Y_{N_{even}}^{1/\sigma}(\tfrac{1}{2} + \tfrac{1}{2}\mu_I^{1-\varepsilon})^{\frac{\varepsilon}{1-\varepsilon}}$$
$$L_{H_{odd}} = Y_{N_{odd}}^{1/\sigma}(\tfrac{1}{2} + \tfrac{1}{2}\mu_I^{1-\varepsilon})^{\frac{\varepsilon}{1-\varepsilon}}$$

where $L_{H_{even}}$ and $L_{H_{odd}}$ is the labor supplied by households in sector A in even and odd periods respectively. For households in sector B, the same expression holds, except that we substitute $L_{L_{even}}$ and $L_{L_{odd}}$ for $L_{H_{even}}$ and $L_{H_{odd}}$, where

$$L_{L_{even}} = Y_{N_{even}}^{1/\sigma}(\tfrac{1}{2} + \tfrac{1}{2}\mu_I^{\varepsilon-1})^{\frac{\varepsilon}{1-\varepsilon}}$$
$$L_{L_{odd}} = Y_{N_{odd}}^{1/\sigma}(\tfrac{1}{2} + \tfrac{1}{2}\mu_I^{\varepsilon-1})^{\frac{\varepsilon}{1-\varepsilon}}$$

The first panel of Table A.1 presents the individual lifetime utility of agents in all states and for different calibrations when excluding real balances from the utility function.[7] The results show that the average individual lifetime utility of agents is lower in the POS than in the CISS and the ZISS. In particular, we observe that $U_{POS_{Average}} < U_{ZISS} \leq U_{CISS_{Average}}$. The disinflation policy reduces social welfare when computed as a simple average of the individual lifetime utility of individuals. The individual lifetime utility of agents is higher in the CISS than in the ZISS. The reason is that output is higher in the CISS.

The second panel of Table A.1 presents the individual lifetime utility of households in all states if I do not exclude real balances from the utility

[7]Obstfeld and Rogoff (1995) argue that utility from real balances is neglectable in welfare comparisons.

Table A.1: The Welfare Effects of the Disinflation Policy for Different Calibrations

Excluding Real Money Balances

Individual Lifetime Utility	Cal A*	Cal B*	Cal C*	Cal D*
CISS of Sector A	-1,50468	-1,07910	-1,73580	-3,29500
CISS of Sector B	-1,50503	-1,07930	-1,73610	-3,29510
CISS avg	-1,50486	-1,07920	-1,73595	-3,29508
ZISS	-1,50517	-1,07920	-1,73600	-3,29510
Post Disinflation Sector A	-1,51988	-1,08820	-1,74370	-3,30030
Post Disinflation Sector B	-1,51337	-1,07720	-1,73488	-3,29510
Post Disinflation avg	-1,51663	-1,08273	-1,73929	-3,29770

Including Real Money Balances

Individual Lifetime Utility	Cal A	Cal B	Cal C	Cal D
CISS of Sector A	2.40120	12.6476	11.4707	23.3531
CISS of Sector B	2.40090	12.6471	11.4705	23.3530
CISS avg	2.40100	12.6474	11.4706	23.3531
ZISS	2.53710	12.7264	11.5159	23.2853
Post Disinflation Sector A	11.2160	17.8407	14.5183	29.9966
Post Disinflation Sector B	11.2225	17.8517	14.5271	30.0019
Post Disinflation avg	11.2193	17.8462	14.5227	29.9993

*Calibrations of Parameter Values as in Table 3.2

function. I now find that $U_{POS_{Average}} > U_{ZISS} > U_{CISS_{Average}}$. The individual lifetime utility of agents is higher in the ZISS than in the CISS. The reason is that real balances are higher in a non-inflationary steady state than in an inflationary one. This appears to more than offset the fact that output is higher in the CISS than in the ZISS. A somewhat surprising finding in Table A.1, however, is that the average individual lifetime utility is higher in the POS than in all other states. The explanation is strongly related to the discussion about the second feasibility constraint. I already discussed the fact that, as the nominal interest rate approaches its lower bound, money demand in odd periods approaches infinity. This implies that odd periods' real balances must become very large in the post-disinflation state as the nominal interest rate moves very close to its lower bound. And the calibrations used in Table A.1 are examples in which the economy is close to violating the feasibility constraint. Moving further away from violating the feasibility constraint, i.e. choosing lower values of β, ε, ζ and μ_I, yields $U_{ZISS} > U_{CISS_{Average}} > U_{POS_{Average}}$.[8] In sum, if we exclude real balances from the utility function, the disinflation policy unambiguously reduces social welfare. If we do not exclude real balances, the disinflation policy may even increase the average individual lifetime utility of households depending on the calibration. It is, however, common to think of the demand for real balances to be subject to a satiation point that is set relative to the demand for consumption goods. The introduction of a satiation point for real balances might remove the above result as it would restrict odd period's money demand to reasonable levels.

[8]A calibration for which this is true, is $\beta = 0.9$, $\varepsilon = 1.4$, $\zeta = 1.2$ and $\mu_I = 1.00001$

A.3 Appendix for Chapter 4

Date	Interest Rate Decision	Bloomberg Median	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	Liffe10 Surprise	Liffe15 Surprise
6/6/97	6.5	6.25	Missing	Yes	No	No	No	No
10/7/97	6.75	6.75	Missing	No	No	No	No	No
7/8/97	7	6.75	Missing	Yes	No	No	No	No
11/9/97	7	7	Missing	No	Yes	No	No	No
9/10/97	7	7	Missing	No	No	No	No	No
6/11/97	7.25	7	Missing	Yes	No	No	Yes	Yes
4/12/97	7.25	7.25	Missing	No	No	No	No	No
8/1/98	7.25	7.25	Missing	No	No	No	No	No
5/2/98	7.25	7.25	Missing	No	No	No	No	No
5/3/98	7.25	7.25	Missing	No	No	No	No	No
9/4/98	7.25	7.25	Missing	No	Yes	No	No	No
7/5/98	7.25	7.25	Missing	No	No	No	No	No
4/6/98	7.5	7.25	Missing	Yes	Yes	Yes	Yes	Yes
9/7/98	7.5	7.5	Missing	No	Yes	No	Yes	No
6/8/98	7.5	7.5	Missing	No	No	No	No	No
10/9/98	7.5	7.5	Missing	No	Yes	No	No	No
8/10/98	7.25	7.5	0.1266	Yes	Yes	Yes	Yes	Yes
5/11/98	6.75	7	0.0589	Yes	Yes	No	No	No
10/12/98	6.25	6.5	0.1447	Yes	No	No	No	No
7/1/99	6	6.25	0.0000	Yes	Yes	Yes	No	No
4/2/99	5.5	5.75	0.1279	Yes	Yes	Yes	Yes	Yes
3/3/99	5.5	5.25	0.1281	No	No	No	Yes	No
8/4/99	5.25	5.25	0.0791	No	No	No	No	No
6/5/99	5.25	5.25	0.1069	No	No	No	No	No
10/6/99	5	5.25	0.1093	Yes	Yes	No	Yes	No
8/7/99	5	5	0.0000	No	No	No	No	No
5/8/99	5	5	0.0000	No	No	No	No	No
8/9/99	5.25	5	0.0000	Yes	Yes	Yes	Yes	Yes
7/10/99	5.25	5.25	0.1008	No	Yes	No	Yes	No
4/11/99	5.5	5.5	0.1065	No	Yes	Yes	Yes	Yes
9/12/99	5.5	5.5	0.0559	No	No	No	No	No
13/1/00	5.75	5.75	Missing	No	No	No	No	No
10/2/00	6	6	0.0736	No	No	No	No	No
9/3/00	6	6	0.0861	No	No	No	No	No
6/4/00	6	6	0.1267	No	No	No	No	No
4/5/00	6	6.25	0.0000	Yes	No	No	No	No
7/6/00	6	6	0.0720	No	No	No	No	No
6/7/00	6	6	0.1146	No	No	No	No	No
3/8/00	6	6	0.1006	No	No	No	No	No
7/9/00	6	6	0.0667	No	No	No	No	No
5/10/00	6	6	0.0000	No	No	No	No	No
9/11/00	6	6	0.0000	No	No	No	No	No
7/12/00	6	6	0.3943	No	No	No	No	No

Table A.2: Monetary Policy Committee Meetings, Interest Decisions and Surprise Measures I

Date	Interest Rate Decision	Bloomberg Median	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	Liffe10 Surprise	Liffe15 Surprise
11/1/01	6	6	0.0828	No	No	No	No	No
8/2/01	5.75	5.75	0.0667	No	No	No	No	No
8/3/01	5.75	5.75	0.1248	No	No	No	No	No
5/4/01	5.5	5.5	0.0490	No	No	No	No	No
10/5/01	5.25	5.25	0.0000	No	No	No	No	No
6/6/01	5.25	5.25	Missing	No	No	No	No	No
5/7/01	5.25	5.25	0.0000	No	No	No	Yes	No
2/8/01	5	5.25	0.0000	Yes	Yes	Yes	Yes	Yes
6/9/01	5	5	0.1308	No	No	No	No	No
18/9/01	4.75	Missing	0.1106	Yes	Yes	No	No	No
4/10/01	4.5	4.5	0.0962	No	No	No	No	No
8/11/01	4	4.25	0.0000	Yes	Yes	Yes	Yes	No
5/12/01	4	4	0.0000	No	No	No	Yes	No
10/1/02	4	4	0.0692	No	No	No	No	No
7/2/02	4	4	0.0000	No	No	No	No	No
7/3/02	4	4	0.0000	No	No	No	No	No
4/4/02	4	4	0.0000	No	No	No	No	No
9/5/02	4	4	0.0730	No	No	No	No	No
6/6/02	4	4	0.0449	No	No	No	No	No
4/7/02	4	4	0.0000	No	No	No	No	No
1/8/02	4	4	0.0000	No	No	No	No	No
5/9/02	4	4	0.0706	No	No	No	No	No
10/10/02	4	4	0.1270	No	No	No	Yes	No
7/11/02	4	4	0.0000	No	No	No	No	No
5/12/02	4	4	0.0000	No	No	No	No	No
9/1/03	4	4	0.0464	No	No	No	No	No
6/2/03	3.75	4	0.0000	Yes	Yes	Yes	Yes	Yes
6/3/03	3.75	3.75	0.0797	No	No	No	No	No
10/4/03	3.75	3.75	0.1259	No	No	No	No	No
8/5/03	3.75	3.75	0.0972	No	No	No	No	No
5/6/03	3.75	3.75	0.1067	No	No	No	No	No
10/7/03	3.5	3.75	0.0000	Yes	Yes	Yes	No	No
7/8/03	3.5	3.5	0.0000	No	No	No	No	No
4/9/03	3.5	3.5	0.0000	No	No	No	No	No
9/10/03	3.5	3.5	0.0373	No	Yes	Yes	No	No
6/11/03	3.75	3.75	0.0386	No	No	No	No	No
9/12/03	3.75	3.75	0.0411	No	No	No	No	No
8/1/04	3.75	3.75	0.0406	No	No	No	No	No
5/2/04	4	4	0.0000	No	No	No	No	No
4/3/04	4	4	0.1239	No	No	No	No	No
8/4/04	4	4	0.0527	No	No	No	No	No
6/5/04	4.25	4.25	0.1256	No	No	No	No	No
10/6/04	4.5	4.5	0.0545	No	No	No	No	No
8/7/04	4.5	4.5	0.0000	No	No	No	No	No
5/8/04	4.75	4.75	0.0377	No	No	No	No	No
9/9/04	4.75	4.75	0.0000	No	No	No	No	No
7/10/04	4.75	4.75	0.0533	No	No	No	No	No
4/11/04	4.75	4.75	0.0000	No	No	No	No	No
9/12/04	4.75	4.75	0.0000	No	No	No	No	No

Table A.3: Monetary Policy Committee Meetings, Interest Decisions and Surprise Measures II

Date	Interest Rate Decision	Bloomberg Median	Forecast Dispersion	Bloomberg Surprise	IB10 Surprise	IB15 Surprise	Liffe10 Surprise	Liffe15 Surprise
12/1/05	4.75	4.75	0.0000	No	No	No	No	No
10/2/05	4.75	4.75	0.0000	No	No	No	No	No
10/3/05	4.75	4.75	0.0000	No	No	No	No	No
7/4/05	4.75	4.75	0.0423	No	No	No	No	No
9/5/05	4.75	4.75	0.0000	No	No	No	No	No
9/6/05	4.75	4.75	0.0652	No	No	No	No	No
7/7/05	4.75	4.75	0.0837	No	No	No	Yes	No
4/8/05	4.5	4.5	0.0000	No	No	No	No	No
8/9/05	4.5	4.5	0.0000	No	No	No	No	No
6/10/05	4.5	4.5	0.0000	No	No	No	No	No
10/11/05	4.5	4.5	0.0000	No	No	No	No	No
8/12/05	4.5	4.5	0.0000	No	No	No	No	No
12/1/06	4.5	4.5	0.0000	No	No	No	No	No
9/2/06	4.5	4.5	0.0000	No	No	No	No	No
9/3/06	4.5	4.5	0.0000	No	No	No	No	No
6/4/06	4.5	4.5	0.0000	No	No	No	No	No
4/5/06	4.5	4.5	0.0000	No	No	No	No	No
8/6/06	4.5	4.5	0.0000	No	No	No	No	No
6/7/06	4.5	4.5	0.0958	No	No	No	No	No
3/8/06	4.75	4.5	0.0000	Yes	Yes	Yes	Yes	Yes
7/9/06	4.75	4.75	0.0000	No	No	No	No	No
5/10/06	4.75	4.75	0.0000	No	No	No	No	No
9/11/06	5	5	0.0000	No	No	No	No	No
7/12/06	5	5	0.0000	No	No	No	No	No
11/1/07	5.25	5	0.0926	Yes	Yes	Yes	Yes	Yes
8/2/07	5.25	5.25	0.0853	No	No	No	No	No
8/3/07	5.25	5.25	0.0857	No	No	No	No	No
5/4/07	5.25	5.25	0.0000	No	No	No	No	No
10/5/07	5.5	5.5	0.0619	No	No	No	No	No
7/6/07	5.5	5.5	0.0809	No	No	No	No	No
5/7/07	5.75	5.75	0.0000	No	No	No	No	No
2/8/07	5.75	5.75	0.0000	No	No	No	No	No
6/9/07	5.75	5.75	0.0323	No	No	No	No	No
4/10/07	5.75	5.75	0.0545	No	No	No	No	No
8/11/07	5.75	5.75	0.1254	No	No	No	No	No
6/12/07	5.5	5.75	0.1010	Yes	No	No	No	No
10/1/08	5.5	5.5	0.0456	No	No	No	No	No
7/2/08	5.25	5.25	0.0323	No	No	No	No	No
6/3/08	5.25	5.25	0.1275	No	No	No	No	No

Table A.4: Monetary Policy Committee Meetings, Interest Decisions and Surprise Measures III

Variable	Coefficient Estimate	p-value
Mean Equation		
Constant	1.02E-004	0.254
AR(1)	0.0345	0.073
AR(2)	-8.42E-003	0.669
AR(3)	-0.0439	0.034
Variance Equation		
Constant	3.36E-007	0.000
ARCH(1)	0.0255	0.000
GARCH(1)	0.961	0.000
Briefing	-2.07E-001	0.020
Day1	-2.70E-001	0.027
Day2	5.50E-001	0.012
Log Likelihood	10628.09	

Table A.5: GARCH Model: Surprise if Change in 3 Months Interbank Rate is Greater Than 9

Variable	Coefficient Estimate	p-value
Mean Equation		
Constant	1.05E-004	0.241
AR(1)	0.035	0.071
AR(2)	-7.98E-003	0.686
AR(3)	-0.0443	0.032
Variance Equation		
Constant	3.38E-007	0.000
ARCH(1)	0.026	0.000
GARCH(1)	0.9602	0.000
Briefing	-2.53E-001	0.045
Day1	-1.66E-001	0.300
Day2	4.34E-001	0.076
Log Likelihood	10625.74	

Table A.6: GARCH Model: Surprise if Change in 3 Months Interbank Rate is Greater Than 14

Variable	Coefficient Estimate	p-value
Mean Equation		
Constant	1.16E-004	0.191
AR(1)	0.0355	0.066
AR(2)	-9.16E-003	0.641
AR(3)	-0.0415	0.041
Variance Equation		
Constant	3.42E-007	0.000
ARCH(1)	0.0256	0.000
GARCH(1)	0.9606	0.000
Briefing	-1.73E-001	0.207
Day1	-2.44E-001	0.063
Day2	4.38E-001	0.037
Log Likelihood	10626.74	

Table A.7: GARCH Model: Surprise if Change in Price of 3 Months Interest Rate Futures is Greater Than 9

Variable	Coefficient Estimate	p-value
Mean Equation		
Constant	1.13E-004	0.209
AR(1)	0.0344	0.079
AR(2)	-7.22E-003	0.715
AR(3)	-0.0435	0.033
Variance Equation		
Constant	3.59E-007	0.000
ARCH(1)	0.0274	0.000
GARCH(1)	0.9577	0.000
Briefing	-3.12E-001	0.013
Day1	-7.47E-002	0.556
Day2	4.94E-001	0.041
Log Likelihood	10625.98	

Table A.8: GARCH Model: Surprise if Change in Price of 3 Months Interest Rate Futures is Greater Than 14

Coefficient Estimates (p-values) of Regime Switching AR(1) Model

$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)
$\mu(2)$	-0.29 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)

Coefficient Estimates (p-values) for Endogenous Transition Probabilities

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	0.74 (0.000)	0.86 (0.000)	0.33 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.003)

LogL=-71084

Table A.9: MS Model: Surprise if Change in 3 Months Interbank Rate is Greater Than 9

Coefficient Estimates (p-values) of Regime Switching AR(1) Model

$\mu(1)$	0.44 (0.000)	$\rho(1)$	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)
$\mu(2)$	-0.28 (0.000)	$\rho(2)$	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)

Coefficient Estimates (p-values) for Endogenous Transition Probabilities

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	0.74 (0.000)	0.51 (0.061)	0.43 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.14 (0.000)			-2.34 (0.000)		-0.86 (0.001)

LogL=-71082

Table A.10: MS Model: Surprise if Change in 3 Months Interbank Rate is Greater Than 14

Coefficient Estimates (p-values) of Regime Switching AR(1) Model					
μ(1)	0.44 (0.000)	ρ(1)	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)
μ(2)	-0.28 (0.000)	ρ(2)	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)

Coefficient Estimates (p-values) for Endogenous Transition Probabilities

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	0.74 (0.000)	0.62 (0.011)	0.39 (0.000)	-0.01 (0.000)	0.31 (0.019)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.000)

LogL=-71082

Table A.11: MS Model: Surprise if Change in Price of 3 Months Interest Rate Futures is Greater Than 9

Coefficient Estimates (p-values) of Regime Switching AR(1) Model					
μ(1)	0.44 (0.000)	ρ(1)	-0.12 (0.000)	$\sigma^2(1)$	6.11 (0.000)
μ(2)	-0.28 (0.000)	ρ(2)	-0.11 (0.000)	$\sigma^2(2)$	1.64 (0.000)

Coefficient Estimates (p-values) for Endogenous Transition Probabilities

	Constant	Dummy 1	Dummy 2	Dummy 3	Dummy 4	Dummy 5
Type of Day		Surprise	MPC	All	MPC	MPC
Time		12:05-13:00	12:05-13:00	12:05-13:45	11:30-11:55	11:15-11:55
P^{11}	0.74 (0.000)	0.67 (0.048)	0.44 (0.000)	-0.01 (0.000)	0.30 (0.020)	
P^{22}	7.15 (0.000)			-2.34 (0.000)		-0.87 (0.001)

LogL=-71082

Table A.12: MS Model: Surprise if Change in Price of 3 Months Interest Rate Futures is Greater Than 14

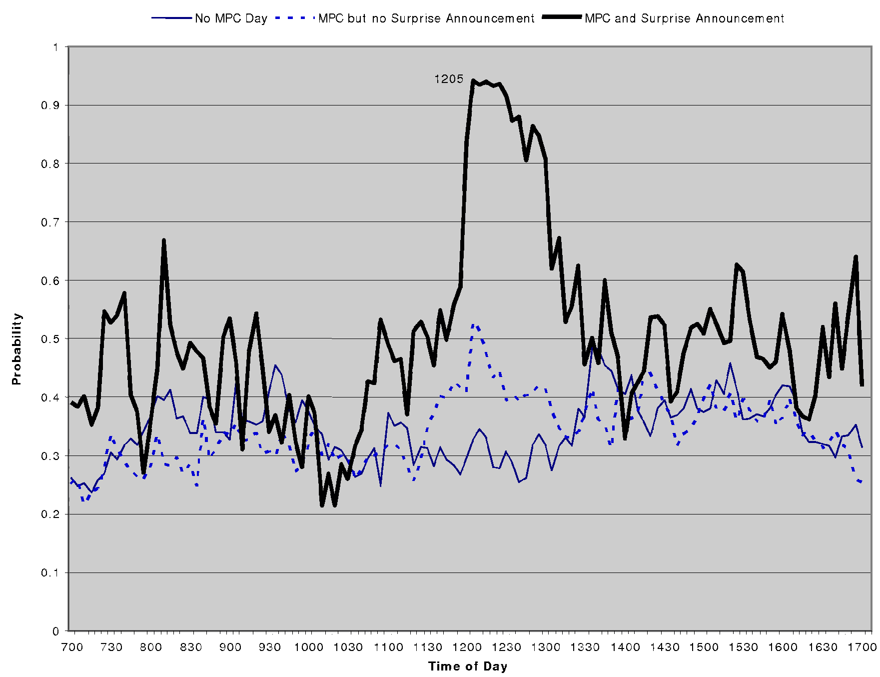


Figure A.3: Smoothed Unconditional Probability of Being in the Informed Trading State - Comparison Between Different Types of Days

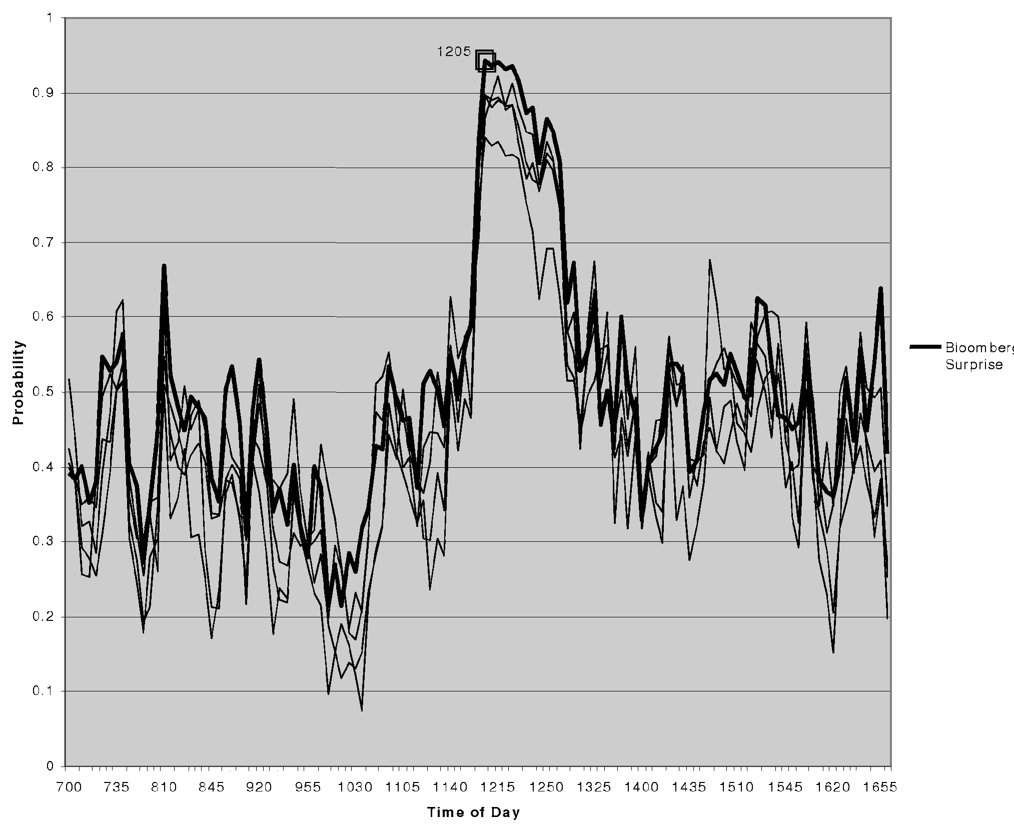
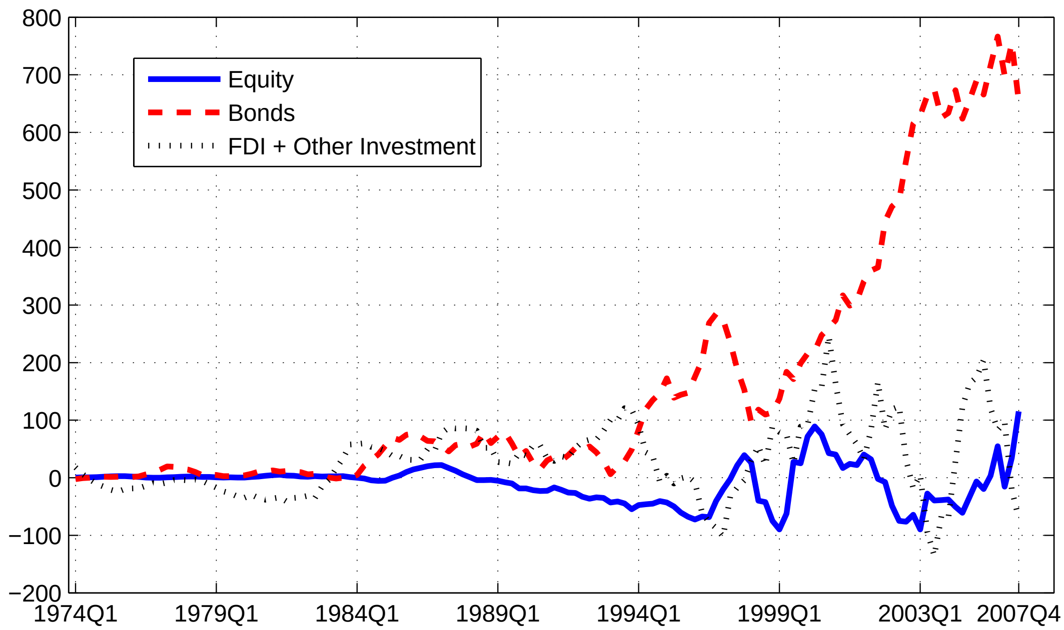


Figure A.4: Smoothed Unconditional Probability of Being in the Informed Trading State on MPC Days with Surprise Announcements - Different Surprise Measures

A.4 Appendix for Chapter 5

Figure A.5: Decomposition of Net Capital Flows to the United States (in billions of USD)



Variable	Definition	Source
Asset Prices	difference between domestic equity returns and foreign equity returns, both measured in US dollars	Bloomberg, mkt indices
Trade Balance	trade balance as a ratio of domestic GDP	IFS
REER	log real effective exchange rate computed using trade weights for a broad set of partner countries	IFS, OECD
Consumption	difference in log private consumption in the domestic and the foreign economy, both measured in US dollars	IFS
Inflation	percentage difference between domestic and foreign CPI inflation	IFS, OECD
Interest rate	percentage difference of short term (money market) interest rates percentage difference of 10 year bond yields	IFS
Reserve ratio	ratio of non-borrowed to total reserves	St. Louis Fed
Financial account	net aggregate inflows of capital in US dollars	IFS
FDI	net inflows of FDI in US dollars	IFS
Other investment	net inflows of other investment in US dollars	IFS
Equity	net inflows of equity investment in US dollars	IFS
Debt	net inflows of debt investment in US dollars	IFS

Table A.13: Variable Sources and Definitions

Figure A.6: Robustness Exercise: Fry and Pagan (2007) Critique

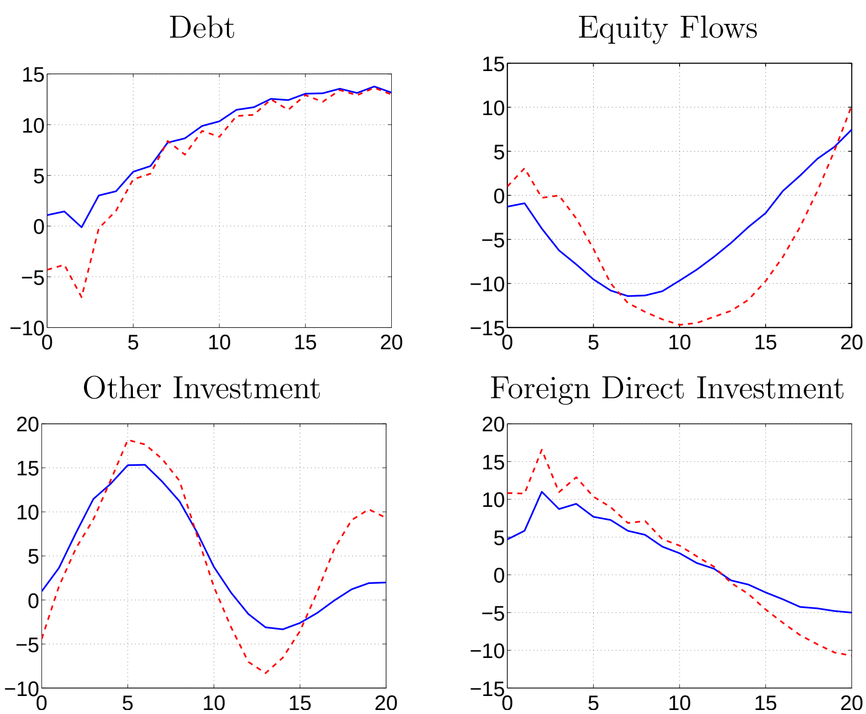


Figure A.7: Equity Parity and Implied Sharpe Ratio

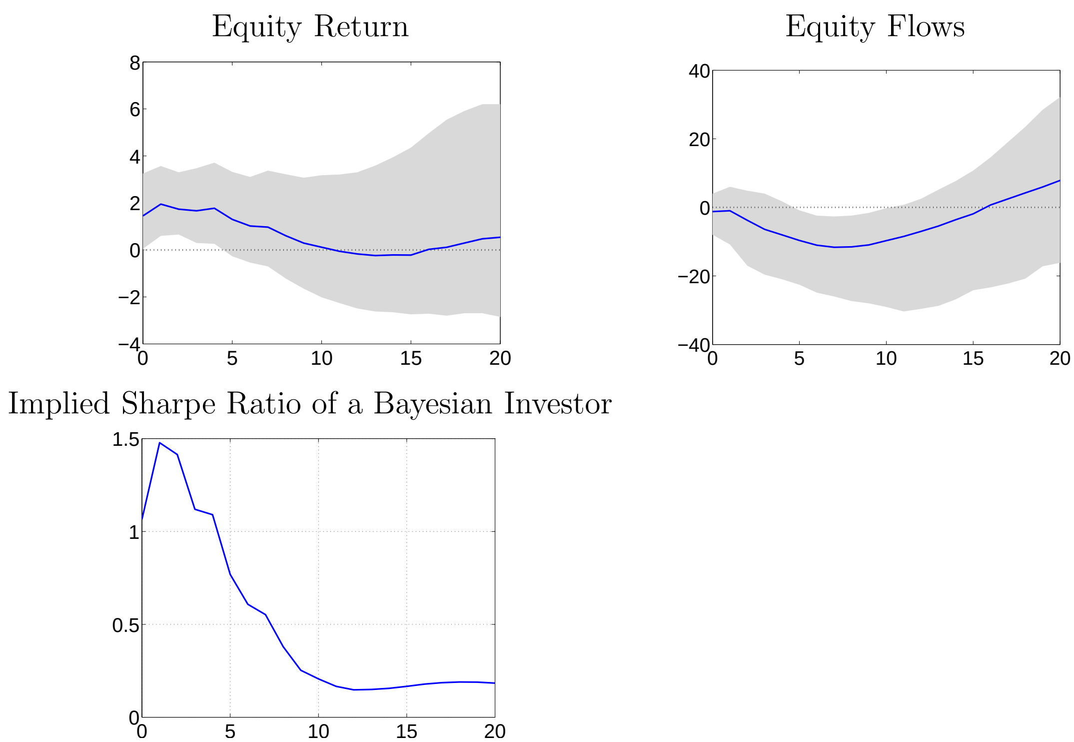


Figure A.8: Robustness Exercise: No Restriction on Consumption

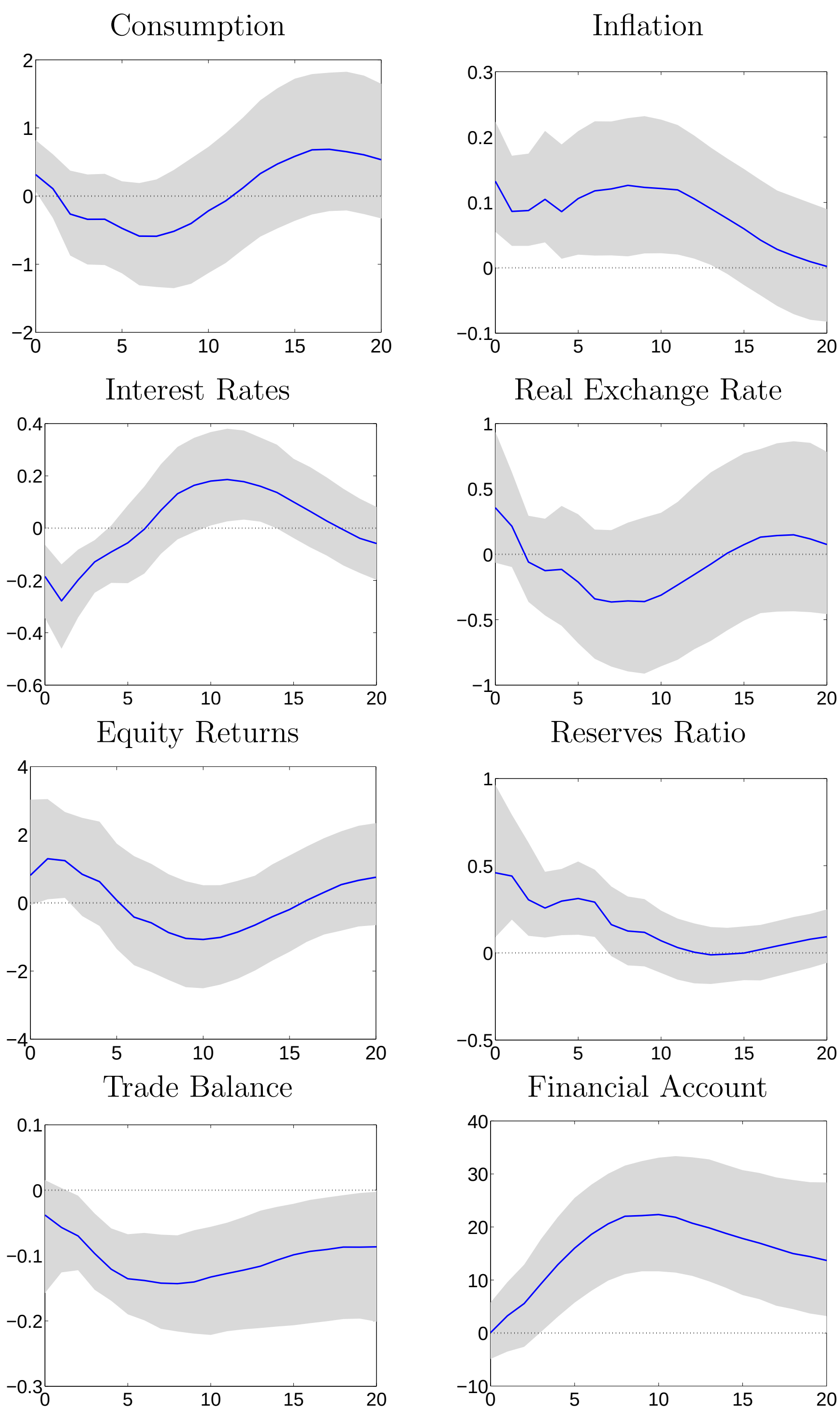


Figure A.9: Robustness Exercise: No Restriction on Consumption - Different Types of Capital Flows

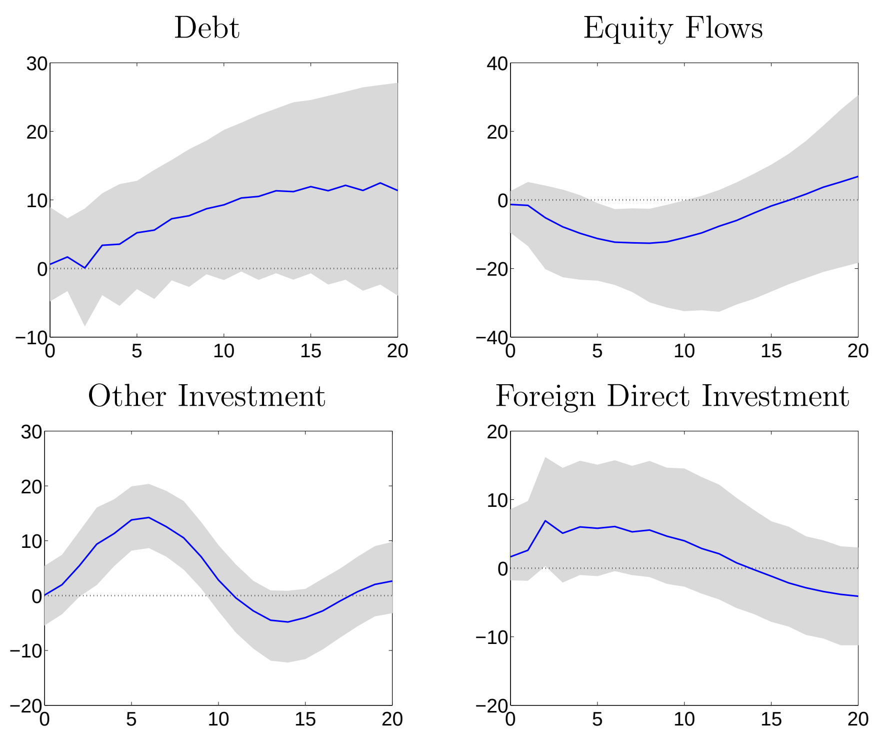


Figure A.10: Robustness Exercise - Identifying Multiple Shocks

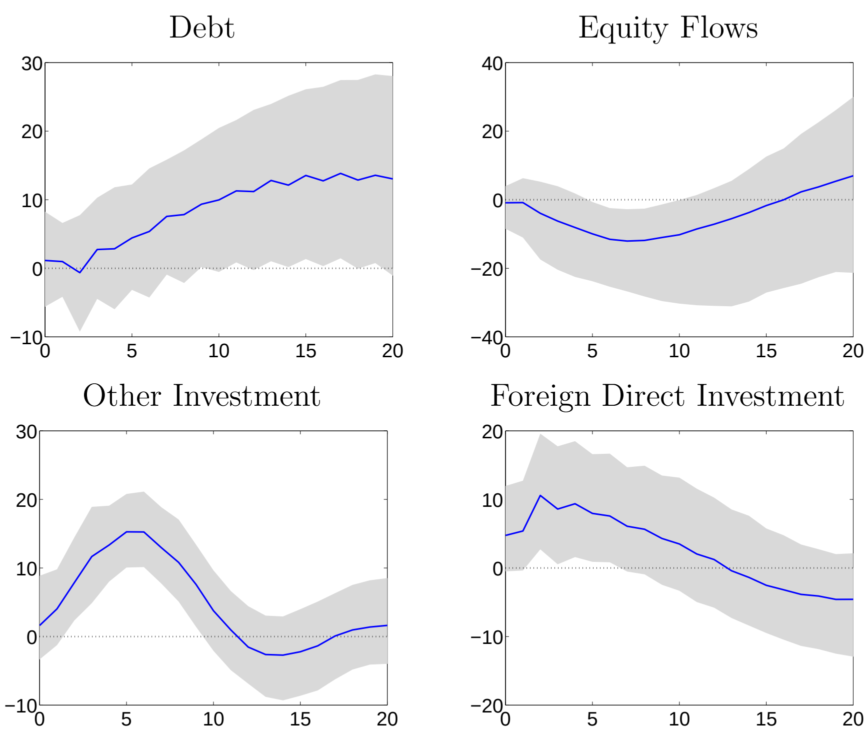


Figure A.11: Robustness Exercise: Equity Prices in Local Currency

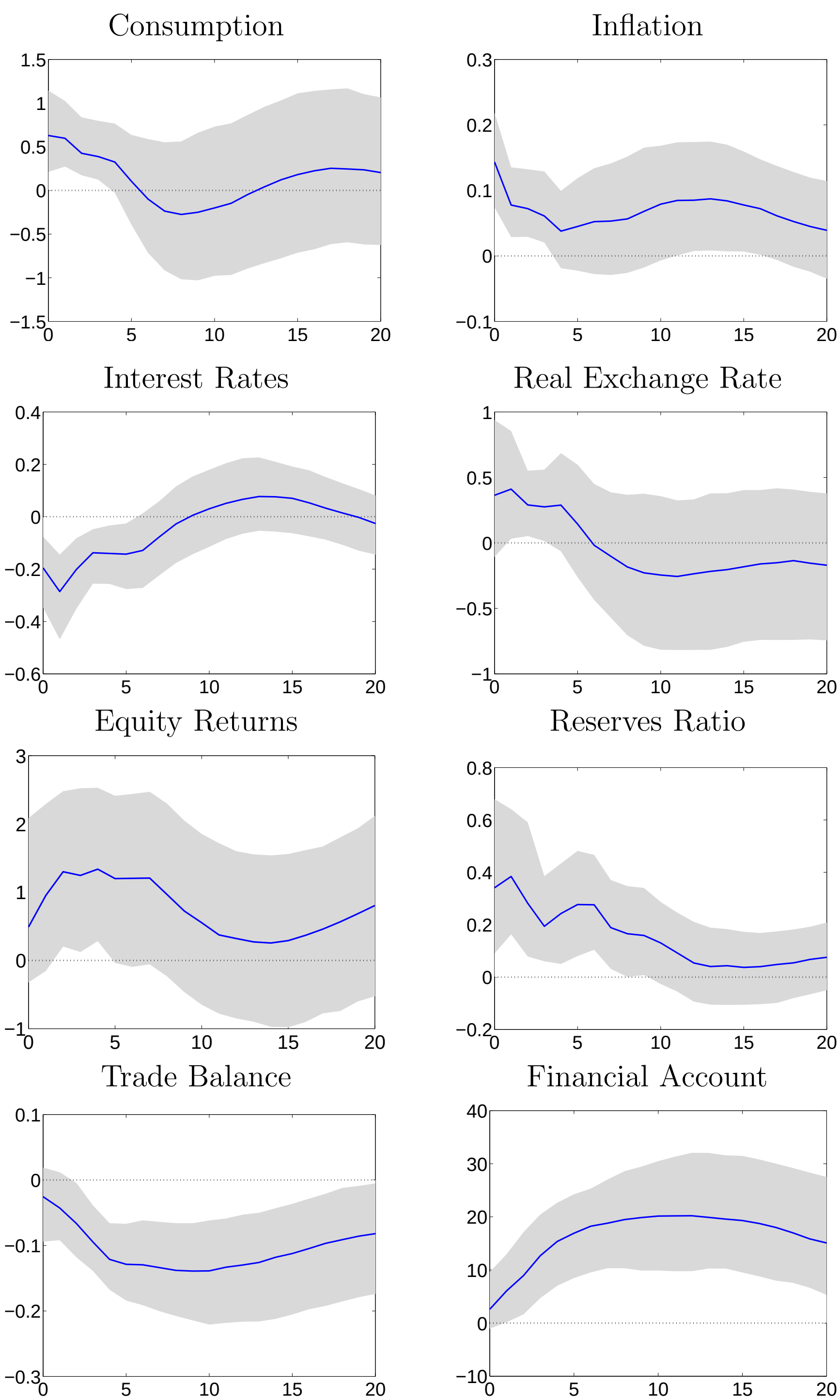


Figure A.12: Robustness Exercise: Equity Prices in Local Currency - Different Types of Capital Flows

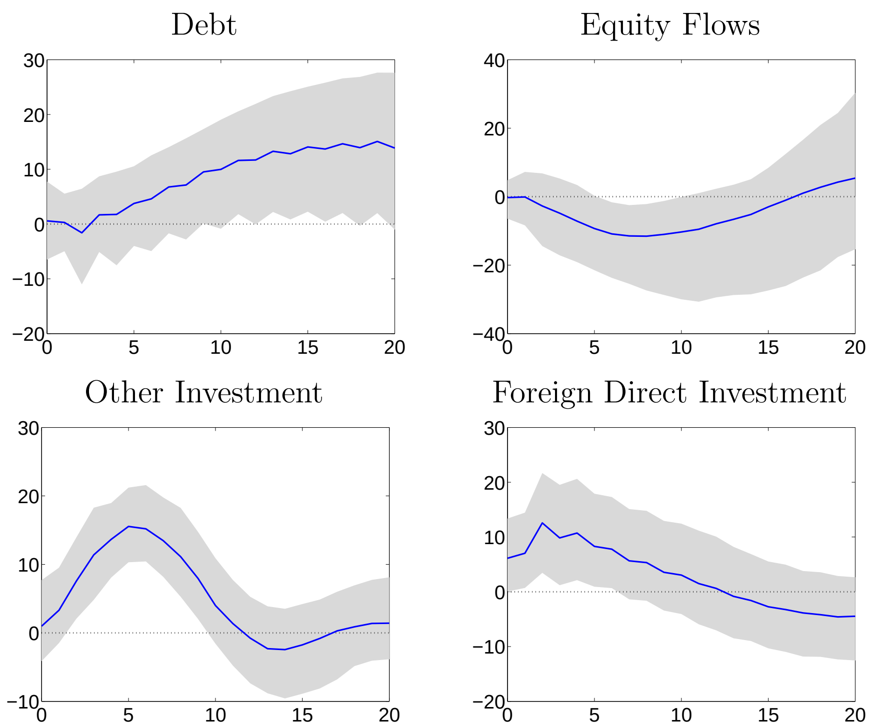


Figure A.13: Robustness Exercise: Capital Flows as a Share of GDP

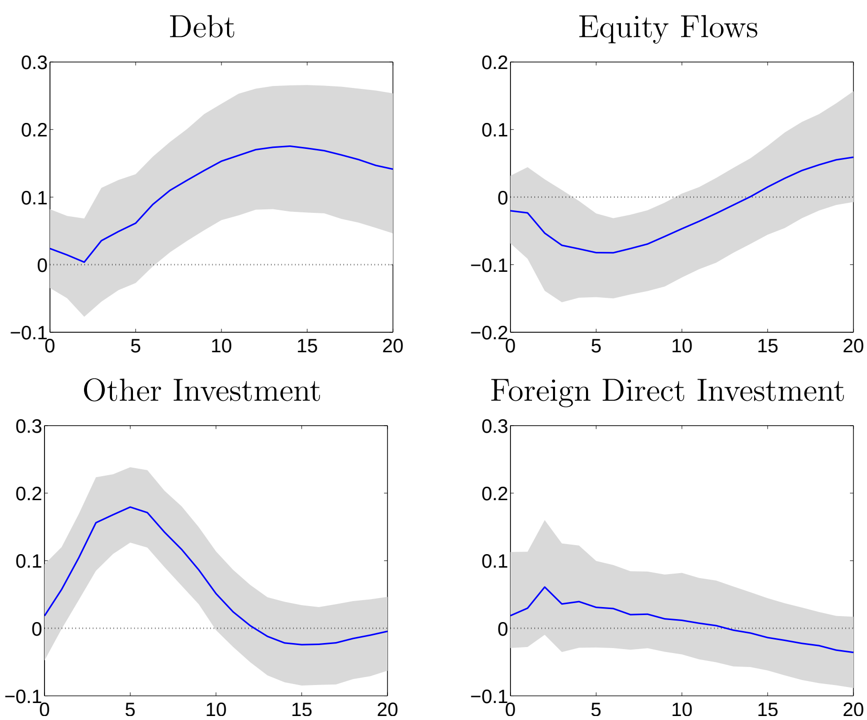


Figure A.14: Robustness Exercise: Including All Types of Capital Flows

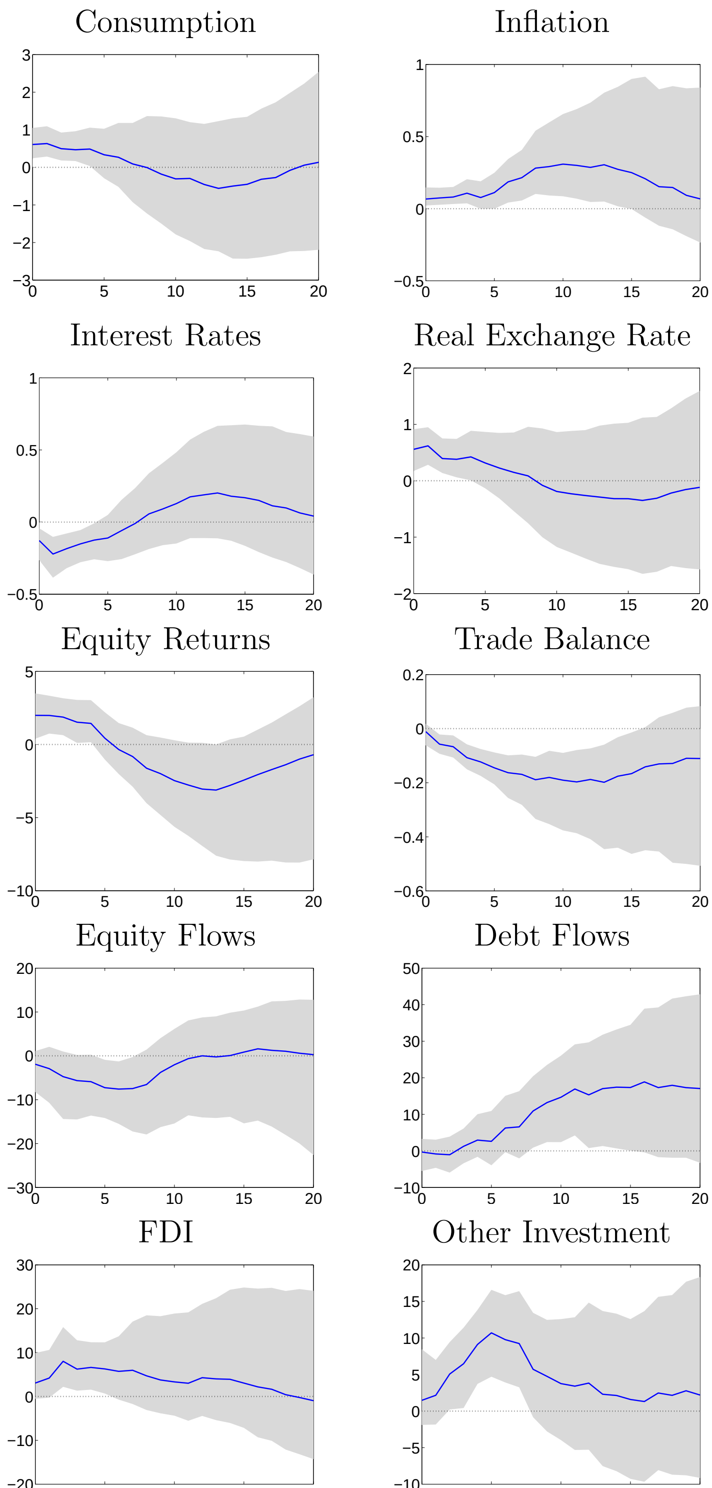


Figure A.15: Robustness Exercise: Extended Rest of the World Sample

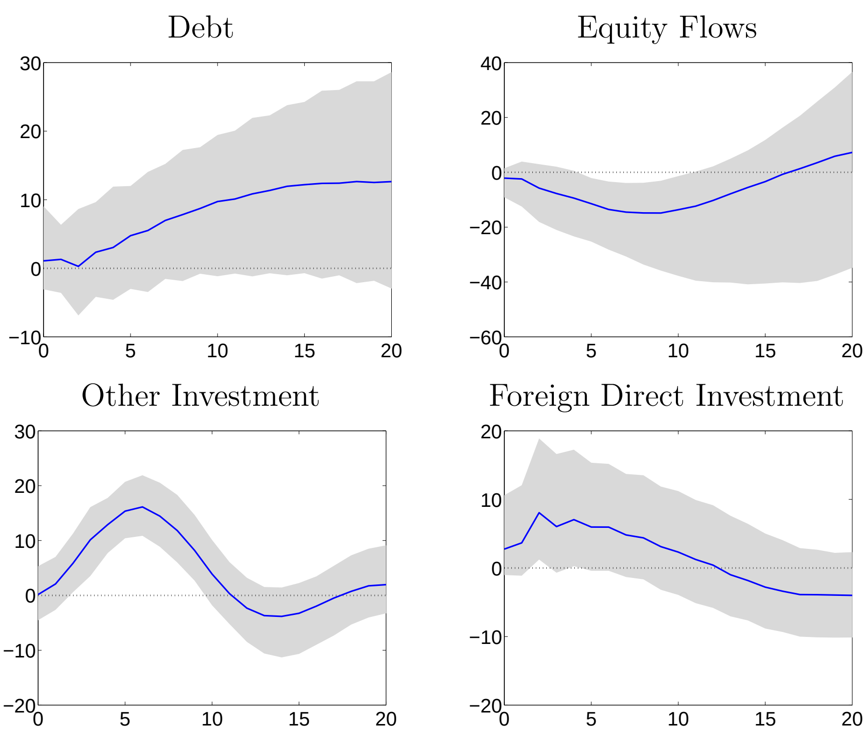


Figure A.16: Robustness Exercise: Including Long-Run Interest Rates