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Three Essays in Empirical International Macroeconomics

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Without my parents' continuous encouragement and support, I probably would not have made it. I dedicate this dissertation to my father; he passed away before I could show it to him.

Chapter 1

Introduction

This dissertation is a collection of three papers in empirical international macroeconomics. All three papers explore a common theme: what are some of the issues that economic policymakers in a developed, small open economy have to deal with? In this respect, the last ten years or so have yielded substantial research progress and new insights in many areas for the (macro)economics profession.

In monetary economics, the introduction of Inflation Targeting (IT) as a monetary policy regime has spurred a great deal of interest, especially from smaller economies — as has, more broadly, cointegration theory from an empirical perspective. Chapter 2 brings these two areas of research together. It analyzes the effects of adopting Inflation Targeting on the monetary policy rule and on money demand for three early inflation targeters, Australia, Canada, and New Zealand, by comparing the pre-IT period with the period after the new framework was introduced.

The main result of Chapter 2 is that the post-IT periods offer less evidence of stable cointegration relationships than the pre-IT periods. In fact, the monetary regime switch might have had profound effects not only on the policy rule but also on money demand, and equilibrium relationships become, therefore, harder to detect in the data. In this sense, the adoption of a new policy regime may have been a preemptive strike to counter the breakdown of a stable money demand. More specifically, the following results emerge. First, the output variable is to a large extent unaffected by the other variables in all three countries, especially the interest rates; in other words, it is often "weakly exogenous." Monetary policy seems to have limited real effects. However, output tends to be "less exogenous" after the introduction of IT. Furthermore, the short-term interest rate becomes exogenous in Canada after the introduction of IT. This surprising finding is interpreted as evidence of increasing monetary integration between Canada and its large neighbor, the United States. Second, the cointegration analysis reveals that, especially in New Zealand and Australia, money demand but also central bank rule-like relationships are more pronounced/stable in the first period than in the second. Third, for New Zealand and Australia, the inflation rate cannot be removed as an "exogenous" variable from the policy rule *before* the introduction of IT. This result is interpreted as proof of the attention that the two reserve banks already devoted to inflation before the actual introduction of the more explicit targeting framework. In Canada, instead, the switch to inflation targeting can be clearly detected in the data: the Monetary Conditions Index plays—as expected—a non-negligible role in the first period, while in the second period the (core) inflation rate cannot be removed from the cointegrating relationship.

In international economics, the surge of the New Open Economy Macroeconomics research agenda is about to provide a new workhorse model of international macroeconomics, taking over from the Mundell-Fleming framework. Based on sophisticated theoretical models rich in empirical implications, this literature also describes the behavior of the current account and its reactions to monetary/nominal shocks—the Dornbusch experiment—as one of its main points of interest. However the empirical literature has been slow to catch up with theoretical developments. In Chapter 3, I try to close parts of this gap by extending the empirical analysis of the small open economy case in the literature to a broader sample of OECD countries, paying particular attention to the G-7 economies.

The main result of Chapter 3 is that there is no consistent response of the current account to nominal shocks across the sample of G-7—and OECD—economies. More specifically, the following results emerge from the analysis of the G-7 economies: First, short-run current account imbalances after nominal shocks are pronounced. Second, countries' current accounts are found to react differently to nominal shocks. The current account surplus predicted by classical theory is not robust across countries. While Japan, Italy, and probably the United States reveal a J-curve effect, other countries manifest purely cyclical behavior. Hence, the results obtained in the literature for the United States cannot be confirmed for a broader sample. Third, while the positive effect on the current account (normalized by its standard deviation) is the highest in Canada and Japan, the relative contribution of a nominal shock in explaining current account variance is maximized in France, the United States, and Italy. There is strong evidence of nominal shocks having short-run real effects, but this evidence is heterogenous across countries. Finally, extending the sample to other (non G-7) OECD countries confirms the conclusion reached for the G-7 economies that there is no consistent reaction of the current account to a nominal shock across counties.

Finally, gauging the slack in the economy is crucial to sophisticated policymaking, both in monetary and fiscal matters. In fact, the cyclical position of the economy has recently (re-)gained a substantial amount of attention, not least due to the discussion about monetary policy rules, the formulation of fiscal rules in the European Union, and the global slowdown in the early years of the new millennium. The output gap—which measures the deviation of GDP from its potential—is a frequently used indicator for the cyclical position of the economy. Defined as the difference between actual and unobservable potential output, the output gap is, however, itself an unobserved variable. Moreover, there are numerous ways to calculate potential output, and the corresponding output gap. In the fourth and last chapter, I estimate a set of output gap measures for a small sample of European countries to evaluate whether the output gap is a concept on which economic policymaking can be based (not only in a small economy of course), and what are the major pitfalls in doing so.

The main results of this last chapter are that output gap measures can yield very different outcomes for a given country depending on the method used to determine potential output and that care should be exercised when dealing with output gap measures—and devising policy recommendations based on them. Moreover, there appears to be little *a priori* reason to prefer one measure over another. The evaluation of a simple forecasting model based on the Phillips curve confirms that the output gap is not always a useful measure to gauge domestic inflationary pressures, and that no specific gap measure consistently dominates all other measures or a simple univariate forecast in this sample of European countries.

Chapter 2

The Effects of Inflation Targeting: Evidence from the Early Movers

2.1 Introduction

During the 1990s, a sizable number of central banks around the world modified their monetary policy framework. They moved away from intermediate targets such as a monetary aggregate and focused directly on the ultimate goal of monetary policy, low and stable inflation. This approach is referred to as Inflation Targeting (IT). Ever since the Reserve Bank of New Zealand (RBNZ) led the implementation of this new type of policy framework in 1989, a lively academic discussion has set in, focusing on the theoretical foundations of IT as well as on empirical issues.[1]

It is well known that the 1990s was a decade of rather smoothly declining inflation rates in a large number of countries, including most OECD member states. In fact, the average inflation rate in industrialized countries declined slowly from 5.1 percent

[1] For overviews see Bernanke *et al.* (1999), Leiderman and Svensson (1995), Haldane (1995), and, more recently, Loayza and Soto (2002), Truman (2003), and the July/August 2004 special issue of the Federal Reserve Bank of St. Louis Review. The theoretical foundations of inflation (forecast) targeting have been covered in various contributions by Svensson (1997, 1999). Svensson and Woodford (2004) compare inflation-forecast targeting to alternative instrument rules. The textbook by Woodford (2003) also treats inflation targeting in some detail.

in 1990 to 1.4 percent in 1999.[2] In most countries that decided to adhere to the new regime, instead, the CPI inflation rate showed a *discrete* drop in the year after implementation. A prominent example is Canada (see Figure 2.1), where a substantial drop in the 12-month headline CPI could be observed, from 1991 (5.6 percent) to 1992 (1.5 percent), equaling approximately the total decrease of inflation in industrial countries over the ten-year period.[3] Another way of stating this is that the unweighted average of the annual Canadian inflation rate was above that of the industrial countries in the decade before 1992 (6.0 percent vs. 5.0 percent), and below it afterwards (1.4 percent vs. 2.2 percent).

This drop coincides chronologically with the adoption of the IT framework by the Bank of Canada (BoC) and the Canadian Government on February 26, 1991.

The record of other countries that implemented IT at various times shows that, on average, these countries experienced a drop in inflation compared to non-IT countries like the United States, Germany, and France (see the appendix to this chapter, Figure A2.1). Given the dispersion over time, the drop in inflation does not seem to stem from a particularly recessional environment with respect to the world economy. In fact, IT was adopted at very different positions in the business cycle. In the following, I will focus on the striking Canadian experience, as well as on two additional economies that were among the first countries to adopt IT—New Zealand (the "pioneer"), and Australia—and ask what effect the introduction of IT has had on these economies.[4]

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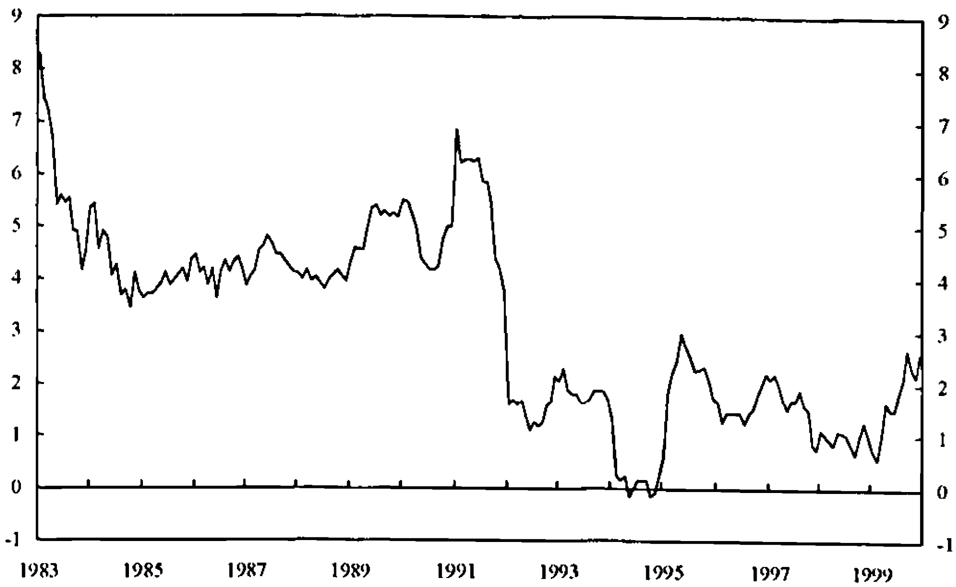
[2]For country-specific data, see Figure A2.1 in the appendix.

[3]Note that the Canadian headline CPI misrepresents "true" inflation in at least two cases: The spike over the first half of 1991, and the consistent trough in 1994. The former is explained with the introduction of a General Sales Tax (GST) in Canada, which was anticipated to have an effect on the CPI change in the order of magnitude of 1.5 to 2 percent; see Bernanke *et al.* (1999). The latter is attributed to a "sharp cut in tobacco taxes designed to curb cigarette smuggling across Canada's border with the United States"; see OECD (1994), p. 25.

[4]During the late 1980s and especially early 1990s, many (OECD) countries adopted a new monetary policy regime together with other important policy measures in the fiscal and structural sphere due to weak macroeconomic performance in the years before the regime change. Selecting such a small sample—while intuitive from an empirical perspective—certainly raises the question whether the results are biased in one way or the other, e.g., toward "greater change." However, the thrust of

Figure 2.1: Canada: 12-month CPI Inflation Rate, 1982-99



Source: International Monetary Fund, *International Financial Statistics (IFS)*.

To investigate the impact of IT, I compare in this chapter cointegrated VAR-models that describe the three economies over roughly symmetric time periods before and after the change in regime. Estimates of long-run relationships and short-run adjustments will provide a picture of the underlying structure of the economy. In particular, I try to identify (possibly changing) relationships that can be interpreted as a monetary policy rule and as a money demand equation to gauge the interaction between the regime change and the underlying economic structures of the three economies. While a change in the policy rule could be interpreted as a consequence of the change in policy framework, the reasoning is somewhat less clear with regard to money demand. The collapse of a stable money demand—*de facto* a decoupling of developments in money supply and prices—is one of the major reasons for many central banks to rely less on the quantity theory of money for policymaking. In that

this chapter is not so much to develop conclusions for a broad set of countries based on this small sample. Instead, the analysis should be viewed as in-depth case studies of the introduction of IT in the three leading implementers.

sense, the adoption of IT could have also preempted a change in shape or a complete breakdown of such an empirical relationship. The analysis is therefore not so much geared towards a normative assessment of IT as having been successful or not, but towards the understanding of "what has changed," along the lines of Juselius (1996).[5]

The remainder of this chapter is organized as follows: Section 2.2 assembles some basic facts related to the policy change in New Zealand, Canada, and Australia and explores briefly the emerging empirical literature on adopting IT as a monetary policy regime. Section 2.3 reviews the main issues in the monetary policy framework adopted in the three countries and section 2.4 provides a brief theoretical background. Section 2.5 delivers the empirical analysis. Section 2.6 concludes, and section 2.7 contains the appendices to this chapter.

2.2 Inflation Targeting—The Facts and the Literature

Building on the extensive theoretical literature praising the advantages of rules over discretion, a key question is whether setting explicit goals for monetary policy also matters empirically. In other words, do central banks with explicit targets experience lower inflation (and, at the same time, no disruptive output variability)? In this connection, Fatás *et al.* (2004) recently investigated the importance of quantitative goals for monetary authorities in a panel of 42 countries for the period 1960-2000. Distinguishing between monetary growth, exchange rate, and inflation targets, they find that having a quantitative (transparent) *de jure* target for the monetary authority (as opposed to "opaque monetary objectives"/discretion) tends to significantly lower inflation and smooth business cycles.[6] Hitting the target de facto has further positive

[5] Juselius (1996) compares the behavior of the Bundesbank in two different periods. For a distinct view of Bundesbank policy, see Clarida and Gertler (1997).

[6] See also Atkeson and Kehoe (2001) for a similar argument.

effects. Among the three target types, Fatás *et al.* (2004) identify inflation targets as bringing the most anti-inflationary impact. The differences between targets are small, however, compared to the presence of any transparent target. With regard to output variability—an often contemplated side effect of bringing down inflation—the sample does not reveal increased volatility, implying that adopting explicit targets comes at little cost-but with large benefits: lower inflation.

In practice, the circumstances under which developed countries and emerging markets adopted IT varied widely. A number of them did not have a specific target for monetary policy before they adopted IT, others had very precise targets, e.g., for the exchange rate. Dissimilarities also exist with respect to the speed of adoption. In general, central banks (and the academic profession) hold diverging views on whether the move from the former regime to targeting inflation in their respective countries can be considered gradual or instantaneous. Especially in those countries that chose IT in the aftermath of a currency crisis (e.g., the United Kingdom), adoption was immediate by nature. While the regime change occurred under different circumstances in the three economies I focus on, monetary policy was rather "opaque" in all of them before they introduced IT.

In New Zealand, the Reserve Bank of New Zealand Act of 1989, which set the stage for the adoption of IT, was a product of the consensus that had emerged gradually within governmental circles, and the central bank in particular, in the years following the 1984 election. In fact, the Act was part of a wide-ranging reform package, including also fiscal, trade, and structural issues. Also, the RBNZ did not have a well-expressed intermediate target during the period before IT was introduced, although it allegedly looked at the major monetary aggregates.[7]

In Canada, no such legislative mandate aimed at implementing the new framework. Despite a three-year campaign to promote price stability starting with the Hanson lecture by then-Governor Crow (1988), the policy shift to inflation targeting

[7]See RBNZ (1985), p. 513.

came with little notice to the public on February 26, 1991. The BoC had suspended M1 as an intermediate target in 1982 and had experimented since then with a monetary conditions index (MCI) as a short-run operational target.

In Australia, the broader aggregate M3 served as an intermediate target until its abandonment in 1985. After that, Australia had no clear-cut monetary policy anchor until early 1993, when the governor of the Reserve Bank of Australia (RBA) projected a numerical target of 2-3 per cent in terms of underlying inflation as a desirable outcome.[8]

Based on the impression that the adoption of the new monetary policy regime in Canada coincided with a drop in the average level of inflation (see above), it is tempting to ask whether this is true for other countries, as well, and whether there is empirical evidence for failure or success of IT. Visual inspection of the data indicates that in most countries the average inflation rate has come down notably since they adopted IT (see Figure A2.1). Empirically, however, the assessment is complicated by concerns about economic policy evaluation à la Lucas (1976), impeding out-of-sample simulation using within-sample estimation if the model's parameters are policy-dependent. Notwithstanding this, an empirical literature evaluating the introduction of IT is emerging.

Early on, this literature has provided a number of interesting case studies; see the contributions by Bernanke *et al.* (1999) and by Mishkin and Posen (1997). In early empirical work, Honda (2000) and Groeneveld *et al.* (1998) do not find convincing evidence for their claims that IT has had a significant impact on macroeconomic variables and that this strategy is superior to intermediate monetary strategies in building monetary policy credibility. Nadal-De Simone (2001) investigates the possible costs of introducing IT by comparing conditional output variance before and after. Consistent with the results in Fatás *et al.* (2004), he finds some evidence that the decline in inflation variance has not been accompanied by an increase in output

[8]See Fraser (1993), p. 2.

variance—with the possible exception of Canada.

More recent contributions to the empirical IT literature include: Choi *et al.* (2003), who find using a Markov-switching model that IT has (i) significantly lowered the volatility of the inflation rate in New Zealand and (ii) led to a structural change in real GDP growth rate; Hu (2003), whose empirical results suggest that inflation targeting does play a beneficial role in improving the performance of inflation and output separately, but who only finds limited support for the proposition that the adoption of inflation targeting improves the trade-off between inflation and output; Ball and Sheridan (2003), who conclude that controlling for regression to the mean in core variables, there is no empirical evidence that IT improves economic performance; and Kuttner (2004), who calls the empirical evidence so far "rather mixed." The July/August 2004 special issue of the Federal Reserve Bank of St. Louis Review contains a number of articles that discuss the success of IT. Broadly speaking, these papers attribute some credit in the deflationary experience in many countries to the introduction of IT. For example, Levin *et al.* (2004) document that in industrialized economies, IT has played a significant role in anchoring long-run inflation expectations by delinking expectations from realized inflation.

So far, cointegration analysis has rarely been used to identify potential effects of IT. This approach, by explicitly considering the long-run aspects, appears to offer an additional point of view almost absent from the empirical IT literature so far. Among the few notable exceptions are: Lee (1999), who finds only scant evidence of regime shift effects in Canada, New Zealand, and the United Kingdom; Karamanou (1999), who compares Taylor rules for the United States, Germany, and the United Kingdom using a variety of empirical techniques (including error correction models); and, to a lesser extent, Valadkhani (2002), who takes into account the IT framework and documents a stable demand for broad money in New Zealand over the period 1988–2002 without explicitly addressing the potential effects related to the introduction of IT, though.

2.3 Inflation Targeting as a Monetary Policy Regime

In this section, the main characteristics of an IT framework are briefly reviewed, outlining the way in which they have been dealt with by the monetary authorities in Canada, Australia, and New Zealand.[9]

Inflation Targeting has been described by the International Monetary Fund as "...the **public declaration of a quantitative target for inflation** in the medium run, coupled with a commitment by the central bank to pursue and reach this target."[10] This straightforward definition is subject to considerable interpretation in a number of directions.

First, the central bank is supposed to be (at least instrument) independent of the fiscal policymaker.[11] The move to increased central bank independence is commonly accompanied by improved accountability (to democratic institutions, e.g., the parliament) and continuous communication between the central bank and the general public in order to foster transparency and understanding of monetary policy issues. All three central banks in this study fulfill this prerequisite, including by publishing various reports and by reporting to their respective parliament.

Second, there should be no conflicting goal of monetary policy, e.g., an additional

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[9]See BoC (various issues), RBA (2004, 1997), and the publications related to Svensson's (2001) review of monetary policy in New Zealand. For an introductory literature overview, see Billmeier (1999), who also discusses issues concerning the central bank's transparency and credibility. More recently, see Svensson and Woodford (2004).

[10]See Blejer *et al.* (*2000*), p. 5 (bold in original). A similar definition is proposed by by Bernanke *et al.* (1999) and accepted by the then-Governor of the Reserve Bank of New Zealand; see Brash (2002). Svensson (1997) goes further to point out that inflation targeting implies in fact inflation *forecast* targeting; that is, the central bank's inflation forecast becomes itself an explicit intermediate target.

[11]For example, the economy should not show any sign of fiscal dominance. Spill-over effects from fiscal to monetary policy might be due to, among other reasons, government borrowing from the central bank, or to an underdeveloped taxation system which relies on seigniorage revenues, see Debelle (1997) and Masson *et al.* (1997). This does *not* imply that the central bank must also set its own goals (goal independence) which can be assigned by law or by the government, see Debelle and Fischer (1994).

exchange rate target, in order to guarantee successful IT.[12] The countries chosen are the only economies that adopted IT at an early stage and at the same time did not form part of an exchange rate arrangement, neither before nor after its introduction. Nevertheless, the exchange rate is of key importance for all three countries.

Third, the monetary authority has the choice among a variety of price level measures. In practice, however, only two versions of the CPI have been targeted. A number of central banks focus on so-called "core inflation," an index which commonly excludes food and energy prices as well as first-round effects of indirect taxation from the CPI. This index is, therefore, less volatile and describes more precisely the effects of monetary policy by excluding important transmission channels from abroad and from the fiscal policy stance. Alternatively, the monetary authority could target the "headline" CPI, which is more easily understood by the general public. In my sample, the BoC has allegedly looked at both rates,[13] while monetary authorities in Australia, as well as in New Zealand (with some delay), aimed predominantly at the underlying rate of inflation. Given the extraordinary effects that blur the picture of the headline inflation rate in Canada (see above, Figure 2.1), the core rate will be used in the empirical analysis that follows, while I will have to employ the non-corrected CPI in Australia and New Zealand for at least two reasons. First, the availability of the time series of underlying inflation (for New Zealand and Australia) is basically restricted to the period after the introduction of IT. Second, the central banks in both countries have chosen to further modify *ad hoc* published time series, such that it is difficult to identify the actual definition of the target variable, also because the composition may have changed over time. Both central banks admit that the headline CPI serves

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[12]This is the case of Spain, which—while adhering to IT—had to respect at the same time the restrictions given by the ERM, i.e., limited exchange rate fluctuations. In fact, adoption of IT has had a insignificant effect on inflation, see Figure A2.1 in the appendix to this chapter.

[13]Even though the official target is the headline inflation rate, strong cases have been made in favor of the core rate, see Dion (1999). This point is of general importance: The Canadian monetary authorities often state (e.g., in Freedman (1995), p. 59) that although there is only a small number of core indicators, the BoC does not limit itself to those but takes an eclectic ("discretionary") view when assessing monetary policy conditions.

at least as a "yardstick against which the Bank should be assessed."[14]

Finally, the target (or more precisely: the ultimate *goal*) has to be specified numerically and over time. In Canada, for example, the first goal was set in February 1991 at a target rate of 3 percent for a 12-month CPI increase by the end of 1992. Since then, the BoC has gradually scaled down the target to a ± 1 percentage points band around a yearly inflation rate of 2 percent over the medium term, i.e., 2–3 years. In the pioneer country, New Zealand, numerical targets for inflation are set in the contract between government and central bank, the Policy Targets Agreement (PTA). In case of breach of the inflation target, the RBNZ governor may be dismissed. These PTAs have distinct duration. While the first PTA lasted for only six months, the latest is to last indefinitely. In the case of New Zealand, these contracts also include a number of "escape clauses;" that is, references to shocks that should not count for assessment, such as changes in indirect taxation.

Table 2.1 presents a selection of instruments, targets, and goals under various monetary policy regimes. The Bundesbank, for example, historically used reserve requirements and, later, central bank interest rates in order to control the medium-run target M3, while the ultimate goal was inflation (π).[15] The other countries of the European exchange rate mechanism—acknowledging the Bundesbank leadership—engaged in the foreign exchange (FX) market and practically targeted the exchange rate to the Deutsche Mark with a view to import low inflation. Other intermediate targets include artificially constructed indicators of the monetary environment, such as monetary conditions indices (MCIs). Alternative (long-run) goals include GDP growth/the output gap—more generally, an employment goal—or maintaining purchasing power parity (PPP) relative to a benchmark country, for example combined with an operational crawling peg target for the exchange rate.

[14] RBNZ (1991), p. 17.

[15] For small open economies with access to international capital markets, often the interest rate spread between short and long rates is viewed as crucial, since the long (bond) rate is determined exogenously by the world interest rate (plus a risk premium).

The last line of the table presents the commonly agreed characteristics of IT. It has been stressed in the literature that thinking of IT as a framework without an intermediate target is an incomplete perception and that the monetary authority's inflation forecast acts as such.[16] Under this regime, instruments of monetary policy—the interest rate(s) controlled by the central bank or the interest rate spread—adjust according to the divergence of the inflation forecast from its long-run goal.[17]

Table 2.1: A Selection of Monetary Policy Regimes

Instruments	*Targets*	*Goals*
Reserve requirements	Money stock	π
Interest rates	Exchange rate	GDP/output gap
FX interventions	MCI	PPP
Interest rate (spread)	(π^f)	(Core) π

Consequently, a reliable model for inflation projections is considered a key element for successful IT, given that the effects of monetary policy need some time to work through the transmission mechanism ("long and variable lags"). In Canada, for example, this forecast was based primarily on the Quarterly Projection Model (QPM) during the period covered by this analysis. However, other factors, like monetary aggregates, are taken into account as well, and the QPM outcome is adjusted accordingly. In these three countries—all small, relatively open economies—the exchange rate and ties to the rest of the world play a particular role when assessing the transmission mechanism of monetary policy and domestic inflation. The Canadian authorities have dealt with this problem in a special way, using an MCI to capture monetary policy effects on aggregate demand stemming from the exchange rate and from the domestic interest rate channel. The MCI indicates the degree of tightness

[16]See Svensson (1997). The BoC confirms this view, see Freedman (2000). Conversely, during the period between 1982, when M1 was dropped as intermediate target, and 1991, when IT was introduced, Canadian monetary policy did not have an explicit intermediate target.

[17]It has also been recognized that—notwithstanding the terminology—the output gap can play an important role in Inflation Targeting, at least in its "flexible" form, see Svensson (1999).

in monetary conditions.[18] Over time, however, the BoC scaled down the use of the MCI because observers and market participants misinterpreted the MCI as a precise short-term target for Canadian monetary policy.[19]

2.4 Steady-State Relations

In the subsequent empirical analysis, I will test a number of theoretical economic relationships for their data congruency. In economic theory, often equilibrium relationships are derived as solutions to systems of (static) equations. Especially in macroeconomics, the introduction of the time dimension provides an important watershed for such equilibrium relations: the empirical analysis can verify whether they hold over time when brought to the data. Hence, empirical analysis has to account for the non-stationarity in most of these time series. The Cointegrated VAR framework yields long-run (steady-state) relationships, consisting of integrated series that "move together" over time. While I briefly explore the possibility of the series being integrated of order 2, or I(2), the basic framework of this chapter, however, will treat them as I(1).[20]

All three economies are "small" and "open;" that is, the effects of foreign variables

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[18] The MCI is constructed as a weighted average (3:1) of the change in a 90-day commercial paper rate (relative to a base period, January 1987) and the percentage change (relative to the same base period) in the exchange rate of the Canadian dollar against a currency basket of six countries, reflecting Canadian trade links, and therefore strongly dominated by the USD (85%). No meaning is attached to the level of the MCI. Given that the MCI is a linear combination of variables that are best modeled as integrated of order one over the period of this analysis, it is highly likely that the MCI itself displays the characteristics of an integrated series.

[19] See Freedman (2000). Another reason for the decline in attention paid to the MCI might be due to the fact that, as has been noted by Eika *et al.* (1996), a meaningful use of an MCI rests on a number of assumptions (e.g., strong and super exogeneity of the MCI, choice of the right variables in the cointegration analysis) that are hardly satisfied.

[20] Theoretical research regarding the I(2) model is still ongoing; see, e.g., Nielsen and Rahbek (2003) and Rahbek *et al.* (1999). For recent applications of the I(2) model, see Juselius (2004) and Kongsted (2003).

are potentially important. I will focus on domestic variables, however.[21] Not accounting for international spill-over effects does not prevent the statistical interpretation of the results. In fact, the cointegration property is invariant to an enlargement of the information set. In other words, an extension to foreign variables would enlarge the system but the fundamental conclusions regarding cointegration relations would persist.[22] Moreover, restricting the analysis to the closed economy case is driven by the methodological approach. Larger VAR systems would entail estimating a larger amount of parameters, and, hence, less significant estimates for a given data span. Since the series' length is limited by the historical context, I will leave the analysis of international spill-over effects for future research.[23]

This chapter focuses on the potential differences between two regimes of monetary policy, and I will pay special attention to the concepts of a central bank policy "rule" and of money demand given that: (i) the empirical evidence generally suggests more than one cointegrating relationship; and (ii) in connection with adopting a new monetary policy regime, money demand and a policy rule appear the most interesting—and empirically relevant–concepts. In the early 1990s, none of the countries in the sample made an explicit policy shift from targeting monetary aggregates to targeting inflation. However, also the adoption of IT after a period of discretion or "opaque monetary policy targets" should be detectable in the data—although less clearly than the switch from one well-defined target to another.[24]

The Central Bank Policy "Rule"

Consistent with Table 2.1, central bank behavior can be described by the way the short-term interest rate i^s reacts to possible right-hand-side variables, including the

[21] In the Canadian case, this is not quite correct since the MCI reflects the effective exchange rate toward the six major trading partners (together with a short term interest rate).

[22] This does not hold for short-term adjustment and the common trends analysis, however.

[23] See, e.g., Juselius and MacDonald (2003), who analyze international parity relationships between Germany and the United States.

[24] See Fatás *et al.* (2004).

inflation rate, monetary aggregates, a measure of output, and possibly the MCI. In accordance with the literature—see, for example, Juselius (1996)—I use the spread between short- and long-term interest rates, $(i^s - i^l)$, as the variable of monetary policy, to capture better the transmission mechanism from short to long rates.[25] Therefore, I will look for a relation of the following form:

$$\left(i^s - i^l\right)_t = \gamma_1(\Delta p_t - \pi^*) + \gamma_2(m-p)_t + \gamma_3 y_t(+\gamma_4 mci_t) + \gamma_5 t + \varepsilon_{1t} \qquad (2.1)$$

where ε_{1t} is a stationary error term, and all variables—with the exception of the interest rates—are in logs. For variables that matter, I expect the coefficient γ_i to be positive; that is, the interest rate spread rises with excessive values of inflation Δp_t (above long-run equilibrium π^*); with excessive growth of real money balances $(m-p)$; and of output y (compared to a simple time trend). The MCI will be included as an indicator variable only in the case of Canada. Note that this specification resembles an "augmented" Taylor rule: apart from excessive inflation and a measure of the output gap, the rule acknowledges the potential importance of the money stock.[26]

The adoption of IT would be expected to impact data patterns that relate to other monetary policy regimes. In Canada, for example, the importance of the MCI—an indicator variable under the previous policy regime—should diminish or even disappear. Moreover, real money balances could have played a (limited) role before the introduction of IT, although the reserve banks of Australia and New Zealand did officially not target a monetary aggregate. After the introduction of IT, one would expect the inflation rate to play a prominent role in the monetary policy rule.

[25] In the cointegration literature, the long-term bond rate has often been found to be important for (small,) open developed economies—but also exogenous, possibly reflecting the world interest level; see, e.g., Johansen and Juselius (1990), who provide evidence for Denmark, and Juselius (1996) for a similar conclusion regarding Germany. Gerlach-Kristen (2003) also finds that long-term interest rates have an explanatory role in the Euro area context, but associates them with inflationary expectations.

[26] Traditionally, Taylor rules are estimated in a stationary environment. See Gerlach-Kristen (2003) for a cointegration analysis of Taylor rules in the Euro area.

Money Demand

The demand for real balances $(m-p)$ is assumed to be the sum of the transactions, precautionary, and speculative demand for money and is given by

$$(m-p)_t = \lambda_1 y_t + \lambda_2 \Delta p_t + \lambda_3 \left(i^s - i^l\right)_t + \lambda_4 t + \varepsilon_{2t}, \qquad (2.2)$$

where ε_{2t} is a stationary error term.[27] I expect the coefficients to be: $\lambda_1 > 0$ (due to the transaction motive, holding of real money balances increases with economic activity); $\lambda_2 < 0$ (the opportunity cost of holding money as opposed to real assets); and $\lambda_3 > 0$ (the opportunity cost of holding monetary assets contained in the definition of the monetary aggregate—here M2—as opposed to longer-term bonds).[28]

The impact of IT on money demand is *a priori* less clear compared to the consequences for the monetary policy rule. There seems to be no strong reason for which the importance of the scale variable y—that is, the transaction motive—should change. With regard to the inflation rate, two lines of reasoning are possible. On the one hand, the importance of inflation in the demand for money could diminish since effective IT delivers a lower rate of inflation, and, hence, agents can disregard the opportunity cost of holding real assets. On the other hand, the credibility and communication effect of the new monetary policy regime could entail heightened public awareness toward the inflation phenomenon and, consequently, strengthen the role of inflation in money demand. Finally, the there is no clear reason why the adoption of IT should affect the opportunity cost of holding (quasi-)money as opposed to bonds as represented by the interest rates and the spread. However, with interest rates—and the spread between them—declining substantially in all three economies in the post-IT compared to the pre-IT period (see Figures A2.2–A2.4 in the appendix to this chapter), it could be expected that this variable becomes less important, or at least less defined, after the

[27]See Goldfeld and Sichel (1990) and Ericsson (1999) on the empirical money demand specifications.

[28]See Coenen and Vega (2001) for a more detailed discussion of money demand parametrization issues.

introduction of IT.

An important prerequisite for both relations to hold is long-run price homogeneity; that is, real money balances are integrated of order 1, $(m-p) \sim I(1)$. Given that both prices and nominal money are possibly I(2), the above formulation implies a non-tested assumption, namely that the time series for money and prices cointegrate from I(2) to I(1), in other words: $\{m, p\}$ are CI(2,1). This can cause some confusion since the two series do not necessarily cointegrate; in other words, a common I(2) trend might still drive real money, and the I(1) framework would be inappropriate for the analysis. In what follows, I assume that the nominal-to-real transformation effectively eliminates possible I(2) trends in the data.[29] To corroborate this assumption, I will determine the (reduced) rank of the system in the I(2) framework to see whether there are remaining I(2) trends in the systems considered.

2.5 Empirical Analysis

This section turns to empirical issues. As a first step, I present the data and briefly discuss problems tied to them. The second step will be to fit VAR models to the data; more precisely, I will present six VAR models, two for each country (before and after the introduction of IT). The third step consists of a cointegration analysis. Restricting the (subperiod) VARs allows identification and comparison of the long-run behavior of the economies in terms of the variables that were introduced in the previous section. The long-run cointegrating relationships and the adjustment coefficients between the two subperiods are compared.

The division into subperiods obviously prompts the question of when to break the series. In the Canadian case, I take a rather pragmatic approach and simply cut out

[29]On the nominal-to-real transformation, see Kongsted (2005), Kongsted and Nielsen (2004), as well as Juselius and Toro (1999), who provide a detailed discussion in the context of an I(2) analysis of the Spanish transmission mechanism. See also Johansen (1992), Juselius (1994), and Rahbek *et al.* (1999).

Table 2.2: Sample Splits

Country	*Period 1*	*Period 2*
Canada	1984:01–1990:12	1992:01–2000:05
Australia	1980:01–1992:04	1993:03–2000:02
New Zealand	1982:01–1989:04	1990:01–2000:02

the year of turbulences (1991), such that roughly eight years of monthly observations remain before and after the break.[30] In the first period, the only obvious central bank target is the MCI, since money targeting (M1) had been abandoned in 1982. For the other two countries, New Zealand and Australia, the analysis is conducted on a quarterly basis because monthly data is not available. Again, the break is assumed to happen when IT is introduced, but no "adjustment period" is introduced. Table 2.2 presents the sample splits.[31]

Splitting the sample into two—with the exception of Canada—adjacent subperiods begs a number of critical considerations and caveats. First, the notorious lags in monetary policy transmission likely imply that also a regime change would take some time to affect not only inflation but also the underlying economic relationships. While this is in principle correct, the cointegration approach does not focus on effects in the months immediately following the regime change. Instead, the second subperiod spans between seven and ten years depending on the country, enough for a regime change to manifest in the data. Moreover, the general public to some extent anticipated a stronger anti-inflationary drive, triggered for example by a lengthy legislative process to prepare for the change (New Zealand), or, as in Canada, by remarks of the then-Governor of the BoC Crow as early as 1988.

This leads straight to the second caveat, namely the difficulty to correctly identify the date when to split the sample. While anticipation may have played a role, I

[30]See Figure 2.1. Juselius (1996) uses a similar approach.

[31]Preliminary assessment of the time series extended throughout end-2003 did not lead to significantly different conclusions.

assume that ultimate clarity about the central banks' objective only came about with the public declaration of the inflation targets in the three sample countries.

A third—notionally similar to but practically distinct from the first—caveat arises regarding the speed of transition. Given that changes in economic aggregates are influenced by behavioral changes of a large number of economic agents, it is highly unlikely that all individual agents will react simultaneously to a change in regime. By implication, a regime change is not abrupt, but implies a transitional period. In this connection, Leyburne and Mizen (1999) use smooth transition analysis to determine endogenously the speed of transition between two different trend paths in prices (all items and underlying CPI) for the same country sample analyzed in this chapter. For the all-items CPIs, they find that 90 percent of the transition from one trend regime in prices to the next takes between about three (Australia) and four (New Zealand) years. Similarly, the transition period for the underlying CPI for Canada (which I employ also in this chapter) amounts to about four years. However, a slightly lower degree of transition, 70 percent, is completed after only 2 – 2.5 years.[32] With this in mind, transitional effects may well impact the end of the first and the beginning of the second sample, but are at the same time only small parts of the subperiods under consideration. In addition, the cointegration approach used in this chapter allows for some flexibility in short-run dynamics without contaminating the long-run analysis. Considering transitional effects more explicitly in the analysis (by either expanding the break period or incorporating a transition function) could provide interesting results but is left for future research.

A similar argument applies with regard to the fourth caveat: the presence of nominal rigidities, which could—arguably—impede fast adjustment of economic agents, for example due to wage contracts or price rigidities. However, all three countries

[32]See Leyburne and Mizen (1999), Table 3. Due to the time series properties of the data, the authors do not estimate the transition period of the all-items CPI for Canada. Extrapolating their results for the underlying CPI, the transition period for the Canadian all-items CPI would be much shorter, presumably less than two years.

dispose of relatively flexible labor markets compared to other OECD countries.[33] Moreover, wage and price setting behavior—if not limited to one-year contracts—is usually staggered, such that a substantial amount of contract adjustment is possible over a short span of time. Therefore, nominal rigidities are not considered to crucially affect the results of the analysis.

2.5.1 Data Description

With the exception of the series on Canadian underlying inflation, the data used throughout this chapter stem from the International Monetary Fund's *International Financial Statistics* (*IFS*) database, CD-ROM version, January 2001, in order to provide a maximum of data coherence.[34] The frequency of the data is monthly for Canada, quarterly for New Zealand and Australia. The CATS in RATS procedure (together with PcGive/PcFiml) were used for the econometric modelling.[35]

The preceding introduction set the stage for a more in-depth discussion of data issues, especially of those time series mentioned above and necessary for the following analysis. For all three countries, similar time series will be used. I will focus on each of the variables in turn, starting with the Consumer Price Index CPI_t. This measure (IFS code ...64...ZF...) was chosen instead of the preferred measure of core inflation (both the RBNZ and the RBA define their target in terms of underlying inflation) for the simple reason that both agencies did not offer a time series long enough to enable empirical assessment. The underlying rate for Canada (starting in 1984) comes from Statistics Canada/BoC, provided through Datastream. As was mentioned above, prices can tend to be described best by I(2) processes. Therefore, the log of the CPI is used in first differences (denominated Δp, or Dp in figures).

[33] See OECD (2003a, 2003b, 2004).

[34] In the appendix, Figures A2.2–A2.4 provide plots of the data in levels and first differences (CPI).

[35] See Hansen and Juselius (1994) and Doornik and Hendry (1997), respectively.

For Canada, the output measure (scale variable) is the industrial production series IP_t (code 15666..CZF...), since measures of GDP are not available at monthly frequency. Using this series is however only a proxy of a broader measure of economic activity/income, such as GDP. One solution would be to disaggregate/distribute the lower frequency GDP data. Given that GDP and industrial production move approximately one-to-one over the whole period, I will use, instead, the monthly series for industrial production, being aware of interpretational restrictions.[36] In the quarterly analysis of New Zealand and Australia, I employ seasonally adjusted real GDP data (line 99BVRZF). The series are computed in (log of) real terms, and denominated y.

For all countries, I employ two interest rates when estimating the cointegrating relationships. The short rate, i^s, reflects a short-term bank rate similar across countries (line 60), while the long rate, i^l, consists of a 10-year bond rate (61...ZF). To obtain a comparable monthly yield, Canadian rates (short/long) originally compiled at a monthly frequency as a yearly return have been modified:

$$i_t^{s/l} = \log\left(1 + rate_t^{s/l}/100\right) * 100/12.$$

For Australia and New Zealand, a corresponding transformation is made to obtain quarterly yields.

Finally, the money stock variable included in the analysis is seasonally unadjusted M2, (the sum of lines 34 (narrow money) and 35 (quasi-money) in the IFS). In principle, the definition of the money stock should correspond to the type of short-term interest rate used in the analysis—the "own rate." However, the literature has not been extremely careful to align these two variables for a number of reasons, including data availability, comparability across countries, statistical properties, and

[36]The hypothesis of cointegration of the quarterly measures of GDP and IP over the whole period in a well-specified VAR (2) was accepted with a p-value of 0.91. This fact motivates the assumption that the measures cointegrate also on a monthly basis. Experiments with interpolation/distribution of GDP from quarterly to monthly observations led to unsatisfactory results.

for robustness tests.[37] In this analysis, choosing M2 instead of M1 is driven by the ease of comparability and its statistical properties. In fact, money enters the analysis in real terms—$\log(\frac{M2}{CPI}) = (m2 - p) = realm2$—which implies the assumption of homogeneity of degree one in prices. In other words, there is no money illusion, individuals demand real balances. Separate tests, not reported here, indicated that this assumption holds "best" in most models for M2.

2.5.2 The Statistical Model

In order to analyze the data, I employ a general cointegrated VAR(k) model, written in error correction form

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \mu_0 + \mu_1 t + \varepsilon_t, \quad t = 1, ..., T, \tag{2.3}$$

where X_t is a p-dimensional autoregressive process, k is the lag length, ε_t is an i.i.d. error with mean zero and variance Ω, $\Pi = \sum_{i=1}^{k} \Pi_i - I$, $\Gamma_i = -\sum_{j=i+1}^{k} \Pi_j$, and D_t contains seasonal and intervention dummies. Under the I(1) hypothesis that $rank(\Pi) = r < p$, the decomposition $\Pi = \alpha\beta'$ holds, where α, β are $p \times r$ matrices of rank r, and $\alpha'_\perp \Gamma \beta_\perp$ has full rank $(p-r)$, where $\alpha_\perp, \beta_\perp$ are the orthogonal complements of α, β, and where $\Gamma = \Sigma_{i=1}^{k-1} \Gamma_i$. The trend is restricted to the cointegrating space, i.e., $\alpha'_\perp \mu_1 = 0$, since quadratic trends are not observed in the data.

The moving average representation of this I(1) process defines the data-generating process for X_t as a function of the errors ε_t, the initial values X_0, and the variables

[37] See, e.g., Ahking (2002), who examines both M1 and M2 aggregates in a cointegration framework for the United States using three-month Treasury bill and commercial paper rates as "own rates." In fact, contributions to the cointegrated money demand literature often use various combinations of monetary aggregates and interest rates. See Knell and Stix (2003) who survey more than 500 individual specifications in their meta-analysis of empirical money demand studies. For Australia, Brouwer *et al.* (1993) survey earlier literature and provide evidence for five different monetary aggregates, ranging from currency to broad money. Atta-Mensah (1995) provides similar evidence for Canada, as does Razzak (2001) for New Zealand.

in D_t. It is given by:

$$X_t = C\left(\sum_{i=1}^{t}(\varepsilon_i + \Phi D_i) + \mu_0 t\right) + C^*(L)(\varepsilon_t + \mu_0 + \mu_1 t + \Phi D_t) + A_0, \qquad (2.4)$$

where $C = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp$. $C^*(L)$ is a finite polynomial in the lag operator L, and A_0 is a function of the initial values.

The cointegrating vectors are estimated by reduced rank regression of ΔX_t on (X_{t-1}, t), corrected for lagged differences and the constant, see Johansen (1996), Theorem 6.2. The vector to be autoregressed is given by $X_t = \left((m2-p), y, \Delta p, i^s, i^l(, mci)\right)'_t$.

2.5.3 The Empirical Model

2.5.3.1 Misspecification Tests

Fitting the models to the data, the number of lags was "tested down" and set to $k = 2$ for all models, quarterly and monthly. Centered seasonal, as well as a few other dummies, were included when necessary to account for outliers.[38] Table 2.3 presents some multivariate misspecification test statistics.[39]

Apart from some first order residual autocorrelation for the second period in Australia and New Zealand, the models seem to be well-specified. Univariate tests report for Canada some problems with ARCH effects for the short term interest rate in period 1, and a rather low R^2 (.19) for i^l in the second period, see Table A2.1 in

[38]The limited number of observations calls for a low number of lags. For Canada, the monthly model, these outliers include D8406, D8412, D8602, D9011, D9209, D9309, D9412, D9810, Ds0002, where D is an impulse, Ds a step dummy at yearmonth. After accounting for these outliers, the monthly models with two lags became acceptable. For Australia (quarterly), there was no need to include dummies. For New Zealand (quarterly), one dummy, D8802 was needed to account for an extreme outlier in the monetary aggregate, see Figure A4 in the appendix to this chapter. This is probably due to a redefinition of the aggregate.

[39]Res. AC(1) and Res. AC (4) are tests for autocorrelation in the residuals, first and fourth order, respectively, while the test for normality is based on a multivariate version of the univariate Shenton-Bowman test, see Hansen and Juselius (1994).

Table 2.3: Multivariate Misspecification Tests

Country	*Test Statistic*	*Period 1* $\chi^2(r)$	p-val.	*Period 2* $\chi^2(r)$	p-val.
Australia	*Res. AC*(1)	18.2(25)	0.83	42.7(25)	0.02
	Res. AC(4)	29.1(25)	0.26	20.2(25)	0.74
	Normality	10.3(10)	0.42	6.8(10)	0.74
New Zealand	*Res. AC*(1)	29.1(25)	0.26	38.1(25)	0.04
	Res. AC(4)	34.1(25)	0.11	27.0(25)	0.36
	Normality	11.9(10)	0.29	8.6(10)	0.57
Canada	*Res. AC*(1)	43.4(36)	0.18	34.1(36)	0.56
	Res. AC(4)	44.0(36)	0.17	23.8(36)	0.94
	Normality	5.6(12)	0.94	14.9(12)	0.25

the appendix. The explanatory power of the system (as measured by R^2) between the two periods rises in Australia and falls in New Zealand and Canada.

2.5.3.2 Rank Determination

As mentioned above, the presence of real money (and eventually real GDP or industrial production) in the analysis calls for particular attention to the possible presence of remaining I(2) trends in the system that did not "cancel out" in the nominal-to-real transformation. To account for this possibility, the trace test for the cointegration rank will be taken from the I(2) framework using the maximum likelihood procedure. Since I include a trend restricted to the cointegrating space, the nonstandard asymptotic distributions derived in Rahbek *et al.* (1999) apply. In all economies, there is no evidence for the presence of a remaining I(2) root; see Tables A2.2–A2.7 in the appendix to this chapter. Generally speaking, there seems to be somewhat "less cointegration" in the data in the second period as compared to the first, independently of the country under observation.[40] Table 2.4 summarizes the results from the rank determination in the I(2) framework.

Given that the I(2) analysis employed here does not allow for dummies, the tables

[40] For example, the hypothesis $r \leq 1$ is only borderline rejected for both Australia and New Zealand in period 2.

Table 2.4: Rank Determination

Country	*Period 1*	*Period 2*
Canada	2	1
Australia	2	2
New Zealand	2	2

could be slightly biased. This is, however, not a major concern in the quarterly models, since there is (almost) no need for deterministic dummies. In the monthly Canadian model, instead, misspecification tests indicated the need to account for a few outliers. Based on the finding that there are no I(2) trends in the system, Table 2.5 presents the trace statistic for Canada in the usual I(1) framework, allowing for dummies.[41]

Table 2.5: I(1) Rank Test Canada

H_0	*Eigenvalue*	*trace*	*trace90*
period 1			
$r = 0$	0.646	**207.9**	110.0
$r \leq 1$	0.443	**112.8**	82.7
$r \leq 2$	0.324	**61.8**	59.0
$r \leq 3$	0.277	38.1	39.1
$r \leq 4$	0.143	16.2	23.0
$r \leq 5$	0.042	3.5	10.6
period 2			
$r = 0$	0.495	**144.94**	110.0
$r \leq 1$	0.261	**92.3**	82.7
$r \leq 2$	0.176	47.4	59.0
$r \leq 3$	0.114	28.2	39.1
$r \leq 4$	0.102	16.1	23.0
$r \leq 5$	0.054	5.4	10.6

From Table 2.5, it can be seen that applying the trace test including dummies has non-negligible effects on the test statistics: in both subperiods, the rank of cointegration seems to rise by one, which is intuitive, since the outliers that were deleted from the single series might have obscured the cointegration property otherwise present in

[41] Trace90 gives the 90% quantile of the likelihood ratio test for the cointegration rank. Rejected hypotheses are in bold.

the data. However, the rejection of $r \leq 2$ in period 1 is only marginal.[42] Complementary evidence can be gained from the roots in the companion matrix, once the cointegration rank is imposed (Table 2.6).

Table 2.6: Roots of Companion Matrix

Country	*Rank*	*Period 1*						*Period 2*					
Canada	3	1	1	1	0.73	0.53	0.32						
	2	1	1	1	1	0.48	0.14	1	1	1	1	0.69	0.25
Australia	2	1	1	1	0.75	0.75		1	1	1	0.81	0.81	
New Zealand	2	1	1	1	0.70	0.70		1	1	1	0.72	0.59	

For Canada, it appears that both possible choices—$r = 3$ and $r = 2$—would be acceptable in the first period, since the next-highest roots do not come close to the unit circle. The unrestricted adjustment coefficients (not reported here) indicate significant error correction in three cointegrating relationships, pointing to $r = 3$. For both New Zealand and Australia, Table 2.6 confirms that $r = 2$ is acceptable for both periods. Consequently, I set $r = 2$ for all models in the following, apart from the first period in Canada, where I assume $r = 3$.

2.5.3.3 Weak Exogeneity

The absence of long-run feedback (i.e., weak exogeneity) for a specific variable offers some preliminary understanding of the driving trends of the system. While I refrain from a more complete common trends analysis for the sake of brevity, it is interesting to test for weak exogeneity, since a weakly exogenous variable can be considered a common I(1) trend.[43]

Where univariate tests indicated more than one weakly exogenous variable, this has been tested jointly as a zero-restriction on the adjustment coefficient, see Table 2.7. The test for Australia, period one, distributed as $\chi^2(6)$ was accepted with a

[42] Moreover, the hypothesis would not be rejected at a 95%-level, with the critical value being 62.6.

[43] See Juselius (1996).

Table 2.7: Univariate Tests for Weak Exogeneity

Country	*Period 1*	*Period 2*
Australia	$y, realm2, i^s$	Δp (, $realm2$)
New Zealand	$\Delta p, y$	$y, realm2$
Canada	$mci, y, realm2$	i^s, i^l

p-value of 0.92. In period 2, joint exogeneity of Δp and $realm2$ was rejected at the 2 percent-level, while the system test of Δp being weakly exogenous gave a test statistic of 2.35 ($\chi^2(2)$), with a p-value of 0.31. For New Zealand, the joint exogeneity hypothesis was accepted with a p-value of 0.12 ($\chi^2(4)$: 7.20) in the first period for Δp and y, and with the same p-value ($\chi^2(4)$: 7.34) for the second period, relative to output and real money, however. In Canada, both interest rates are jointly weakly exogenous in period 2 ($\chi^2(4)$: 3.12, p-value = 0.54). In period 1, the monetary conditions index, industrial production and real money can be considered weakly exogenous, the likelihood ratio test yields $\chi^2(9)$: 13.54, with p-value = 0.14.

This preliminary assessment sheds some light on the driving forces (the common trends) of the systems before and after the imposed break.

First, it is interesting to see how (radically) different they are. For example, in the model describing Canada before the introduction of IT, the trend stemming from the MCI may reflect some foreign influence via the exchange rate, but there are also two domestic stochastic trends, one from the real sector (y, or industrial production), and a monetary trend, $realm2$. In the second period, this finding is completely inverted and both interest rates become exogenous. In the Canadian context, this could be interpreted as a growing economic and especially financial integration, such that the U.S. interest rates are also dictating the Canadian financial markets.

Second, the output measure is weakly exogenous in four out of six models (New Zealand P1+P2, Australia P1, Canada P1)—an indicator of the "non-affectedness" of output by the other variables in the system, or the limited real effects of monetary policy in the long run. In other words, output responds in equilibrium more to the variables included in the system after the introduction of IT in both Australia and

Canada.

Third, the price variable ceases to be a common trend in New Zealand after the introduction of IT, but becomes one in Australia. Since both short- and long-term interest rates are included in the system, this could mean that monetary policy (as expressed by movements in interest rates) became less capable of influencing price developments in Australia—a rather worrisome conclusion.

2.5.4 Long-run Identification

In this section, I describe identified cointegrating relationships for the six models, preferably money demand and a central bank rule where possible. No restrictions on the α-vectors (i.e., no weak exogeneity assumptions) have been imposed. The corresponding tables in the appendix to the chapter (A2.8–A2.10) report the $\beta-$vectors in the upper part (standard errors in parentheses), and the adjustment coefficients in the lower part (t-values in parentheses).

2.5.4.1 Australia

The Australian data reveal a money demand relationship $(\widehat{\beta}_1)$with expected signs in period 1 (see Table A2.8). Real money balances are borderline error-correcting, the other variables do not significantly adjust to the long-run equilibrium. This relationship does not hold up in period 2, where the coefficients to the interest rates have the wrong sign and where "money demand" does not respond to the change in the price level any more. Furthermore, it seems that the significant adjustment of the short interest rate and the large adjustment coefficient to Δp (together with significant error correction of real money balances) point to another relationship, which could not be identified.

A similar picture emerges for the second relationship, the central bank rule. In the first period the interest rate spread $(i^s - i^l)$ moves together with the inflation rate and real money, with significant adjustment of the short-term interest rate.[44] No such relationship can be detected after the introduction of IT. The only meaningful result found in the data is given by $\widehat{\beta}_2$, and could be interpreted as an IS curve (inverse output-interest rate relationship), with significant adjustment in the correct direction. A similar relationship was not present in the first period. The size of the coefficients (especially the adjustment coefficients, but also the cointegrating coefficient to i^l) cast some doubts on the latter result. However, the joint hypotheses on the $\beta-$coefficients were comfortably accepted in both periods at the 50 percent-level.

Somewhat surprisingly, from the whole range of possible formulations for the central bank rule in the first period (i.e., excluding one or more of the RHS variables, $(m2-p), y, \Delta p)$ from the cointegrating space, only two are viable, namely excluding y or $(m2-p)$—but not both at the same time. All others (especially the exclusion of the inflation rate) are either rejected or lead to implausible coefficients. In the second period, as described above, no formulation is accepted at conventional levels. This could be interpreted as evidence of the fact that the RBA did indeed take strongly into account inflation when setting the short-term interest rate already before the targeting of inflation was announced. Summing up, there is only a weak case for a second cointegrating relationship, see also Table A2.7.

2.5.4.2 New Zealand

The New Zealand analysis proposes a rather similar picture (Table A2.9). The $\widehat{\beta}_1-$vector can be interpreted as a money demand relationship, with conventional coefficients and error-correcting real balances, while in the second period correct signs can only be achieved if the coefficient to the long interest rate is set to zero; that is, real balances do not respond to the opportunity cost of holding bonds.

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[44]Note that the coefficient to inflation is not significantly different from 1. It has been shown in the context of New Keynesian sticky price models that a coefficient greater than 1 is a sufficient (but not necessary) condition to fulfill the "Taylor principle," i.e., to avoid indeterminacy of the rational-expectations equilibrium price level, see Woodford (2001) and Taylor (1999). In his seminal contribution, Taylor (1993) estimated the inflation coefficient at 1.5.

The second relationship, the central bank rule, shows the expected signs and meaningful coefficients in period 1. Note that the coefficient to inflation is significantly below unity, implying a "passive" behavior of the central bank with respect to inflation. The monetary authority seems to react strongly to (inflationary) output growth. Imposing the interest rate spread as proxy for the instrument of monetary policy in period 2 does not yield meaningful results.[45] Therefore, the bond rate is not included in the cointegrating space of $\widehat{\beta}_2$ in the second period. The vector reflects a relationship between the strongly equilibrium-correcting short-term interest rate and detrended real M2. Just as in the Australian case, also in New Zealand the inflation rate cannot be removed from the cointegrating space *before* the regime switch without provoking a rejection or implausible coefficients. In the second period, however, this is possible. The joint hypotheses on the β–coefficients are borderline accepted in the first period, but comfortably in the second.

2.5.4.3 Canada

As indicated above, three cointegrating relationships are assumed in the first period. $\widehat{\beta}_1$ mimics again the money demand relation, which is rather similar in both periods (Table A2.10). Note that the coefficient to Δp—the opportunity cost of holding money instead of real assets—has roughly tripled in the second period. This could be viewed as evidence of agents' increased awareness with regard to inflation, possibly triggered by the adoption of the IT framework. An interpretational problem arises from the fact that real balances are not significantly equilibrium-correcting in the first period, but do adjust in the second. The inflation rate, in turn, is strongly adjusting in the first period but not in the second.

The policy rule ($\widehat{\beta}_2$) offers two interesting conclusions. First, the short-term interest rate strongly adjusts to disequilibria in the first, but not in the second period.

[45] In fact, the univariate test statistic for long-run exclusion of the bond rate (not reported here) is non-significant for all choices of the cointegration rank, i.e., the bond rate is not helpful in explaining the information present in the VAR system.

This confirms the above analysis, where both interest rates were found to be weakly exogenous after the introduction of IT. This could be due to the increasing integratedness of North-American (financial) markets, with a common interest rate determined in the U.S. market. Second, CPI inflation can be eliminated from the cointegration space in the first period but not the MCI, and vice versa for the second period. This is what one could intuitively expect from the described change in the monetary policy regime (from targeting "nothing," eventually the MCI, to targeting inflation). In the given specification, the interest rate spread rises with the MCI in the first period and with real money and inflation in the second. Note that the coefficient to inflation implies an "active" central bank behavior.

The third cointegrating vector in period 1 presents a relationship visible in the data, the co-movement of the MCI and the bond rate (see Figure A2.4 in the appendix). This comes as no surprise, given the preponderance of an (admittedly shorter-term) interest rate over the exchange rate in the construction of the MCI.[46]

2.6 Conclusion

This chapter has provided an empirical assessment of the effects provoked by the introduction of a new framework for monetary policy in Australia, New Zealand and Canada using cointegration theory. The chapter identified long-run relationships interpretable as monetary policy rules and money demand functions as implied by the data. A number of interesting results emerge from the above analysis.

First, regarding weak exogeneity, the output variable is to a large extent unaffected by the other variables in the system, especially the interest rates. Monetary policy seems to have limited real effects. However, output is "less exogenous" after the introduction of IT. In Canada, the short-term interest rate becomes exogenous after the introduction of IT. This surprising finding is interpreted as evidence of increasing

[46] See footnote 18.

monetary integration between Canada and its large neighbor, the United States. The overnight money market rate becomes a common trend of the system—potentially driven by the U.S. financial market—instead of the outcome of domestic policy.

Second, the cointegration analysis reveals that especially in New Zealand and Australia, money demand but also central bank rule-like relationships are more pronounced/stable in the first period than in the second. After the introduction of IT, an interest rate rule cannot be detected at all in Australia, while in New Zealand, the long-term interest rate does not play any role. In both countries, however, the second cointegrating relation is only borderline stationary.

Third, for New Zealand and Australia, the inflation rate can be removed from the cointegrating space *after* the introduction of IT, but not so *before*. One factor that might contribute to this rather counter-intuitive result is that both central banks after the regime switch in fact targeted core inflation while the time series employed in the analysis refer to the headline CPI due to data availability problems. The implied discrepancy between core and headline CPI is not intuitive, however. In Canada, on the other hand, the switch to inflation targeting can be clearly detected in the data: the MCI plays—as expected—a non-negligible role in the first period, while in the second period removing the (core) inflation rate from the cointegrating relationship is rejected on statistical grounds.

Fourth, the data provide less evidence of cointegration after the introduction of IT in a general sense. In fact, the monetary regime switch might have had profound effects not only on the policy rule but also on money demand, and equilibrium relationships are therefore hard to detect in the data.[47] In this sense, the adoption of a new policy regime may have been a preemptive strike to counter the further

[47] Clearly, the above analysis (not so much for Canada) is also restricted by the rather limited number of observations. This is partly due to the fact that New Zealand and Australia do not collect/publish data on inflation (and industrial production) on a monthly basis. Svensson (2001) harshly criticized this as "striking," (p. 4), adding that monthly startistics are "required to bring data quality up to international standards" (p. 48). On the other hand, a preliminary assassment of data sets extended thoughout end-2003 did not appear to provide significantly different conclusions.

disintegration of a stable money demand, a cornerstone of many other monetary policy frameworks. From a chronological point of view, financial innovation that could have triggered money demand instability preceded the introduction of IT by some 10 years.[48]

[48]See, for example, Freedman (1983).

2.7 Appendices

2.7.1 Tables

Table A2.1: Univariate Misspecification Tests

Country	*Series*	*Period 1*			*Period 2*		
		ARCH(2)	Norm	R^2	ARCH(2)	Norm	R^2
Australia	Δp	1.906	0.760	0.723	0.512	0.486	0.611
	y	7.632	4.692	0.380	0.646	0.441	0.707
	$(m2-p)$	0.065	5.257	0.490	0.411	0.226	0.769
	i^s	6.910	1.088	0.414	2.057	0.996	0.891
	i^l	0.948	3.403	0.414	4.718	3.311	0.781
New Zealand	Δp	2.902	5.047	0.757	2.328	2.028	0.752
	y	0.627	1.549	0.564	1.373	3.938	0.379
	$(m2-p)$	2.115	1.189	0.819	0.589	0.872	0.501
	i^s	1.108	0.190	0.808	1.790	0.351	0.641
	i^l	0.194	2.081	0.636	1.180	2.618	0.522
Canada	Δp	0.949	0.429	0.854	0.413	1.279	0.734
	mci	2.607	3.828	0.563	0.084	1.730	0.567
	y	1.971	0.477	0.683	1.064	3.150	0.410
	$(m2-p)$	2.493	0.401	0.857	1.001	0.745	0.523
	i^s	11.493	1.444	0.593	0.991	4.882	0.530
	i^l	1.686	1.130	0.562	1.476	0.066	0.192

Table A2.2: I(2) Trace Test Canada, Period 1 (P1)

p-r	**r**	$S^{\infty}_{r,s}$						$S^{\infty}_{r,p-r}$
6	0	**1936.9**	**523.2**	**417.5**	**326.7**	**251.9**	**194.3**	**168.5**
		(269.2)	(233.8)	(202.8)	(174.9)	(151.3)	(130.9)	(115.4)
5	1		**454.3**	**348.6**	**257.8**	**181.2**	**123.7**	**98.33**
			(198.2)	(167.9)	(142.2)	(119.8)	(101.5)	(87.2)
4	2			**311.25**	**228.8**	**152.2**	**88.9**	60.4
				(137.0)	(113.0)	(92.2)	(75.3)	(62.8)
3	3				140.36	81.9	64.5	29.4
					(86.7)	(68.2)	(53.2)	(42.7)
2	4					57.2	25.3	12.9
						(47.6)	(34.4)	(25.4)
1	5						2.62	2.17
							(19.9)	(12.5)
p-r-s		6	5	4	3	2	1	0

Note: The number of I(1) components is s and the number of I(2) components is $p-r-s$. Simulated 95 percent critical values in parenthesis. A test statistic higher than the critical value means that the hypothesis is rejected. The table is to be read from upper left to lower right, line by line. Rejected hypotheses are printed in **bold**.

Table A2.3: I(2) Trace Test Canada, Period 2 (P2)

p-r	**r**	$S^{\infty}_{r,s}$						$S^{\infty}_{r,p-r}$
6	0	**665.8**	**513.3**	**399.6**	**304.7**	**228.2**	**161.3**	**128.2**
		(269.2)	(233.8)	(202.8)	(174.9)	(151.3)	(130.9)	(115.4)
5	1		**450.6**	**340.3**	**246.5**	**171.6**	**105.1**	73.1
			(198.2)	(167.9)	(142.2)	(119.8)	(101.5)	(87.2)
4	2			287.4	206.5	139.0	73.0	42.9
				(137.0)	(113.0)	(92.2)	(75.3)	(62.8)
3	3				170.4	100.4	53.0	24.2
					(86.7)	(68.2)	(53.2)	(42.7)
2	4					104.0	39.4	12.5
						(47.6)	(34.4)	(25.4)
1	5						63.7	3.42
							(19.9)	(12.5)
p-r-s		6	5	4	3	2	1	0

Table A2.4: I(2) Trace Test New Zealand, P1

p-r	**r**			$S^{\infty}_{r,s}$			$S^{\infty}_{r,p-r}$
5	0	**718.0**	**232.0**	**194.9**	**169.7**	**147.8**	**131.5**
		(198.2)	(167.9)	(142.2)	(119.8)	(101.5)	(87.2)
4	1		**252.8**	**148.6**	**121.9**	**98.1**	**81.8**
			(137.0)	(113.0)	(92.2)	(75.3)	(62.8)
3	2			**99.7**	**74.4**	**54.7**	41.5
				(86.7)	(68.2)	(53.2)	(42.7)
2	3				45.7	26.0	12.9
					(47.6)	(34.4)	(25.4)
1	4					16.9	3.4
						(19.9)	(12.5)
p-r-s		5	4	3	2	1	0

Table A2.5: I(2) Trace Test New Zealand, P2

p-r	**r**			$S^{\infty}_{r,s}$			$S^{\infty}_{r,p-r}$
5	0	**294.5**	**224.6**	**182.8**	**151.3**	**124.6**	**113.6**
		(198.2)	(167.9)	(142.2)	(119.8)	(101.5)	(87.2)
4	1		**195.6**	**146.9**	**110.8**	**84.8**	**63.7**
			(137.0)	(113.0)	(92.2)	(75.3)	(62.8)
3	2			**117.4**	**79.9**	**57.0**	36.7
				(86.7)	(68.2)	(53.2)	(42.7)
2	3				63.3	31.0	14.7
					(47.6)	(34.4)	(25.4)
1	4					34.4	4.3
						(19.9)	(12.5)
p-r-s		5	4	3	2	1	0

Table A2.6: I(2) Trace Test Australia, P1

p-r	r	$S^{\infty}_{r,s}$					$S^{\infty}_{r,p-r}$
5	0	**1076.2**	**222.4**	**174.9**	**140.3**	**111.9**	**101.0**
		(198.2)	(167.9)	(142.2)	(119.8)	(101.5)	(87.2)
4	1		**186.4**	**138.0**	**103.4**	**79.1**	**65.1**
			(137.0)	(113.0)	(92.2)	(75.3)	(62.8)
3	2			**112.4**	**78.0**	**56.1**	37.9
				(86.7)	(68.2)	(53.2)	(42.7)
2	3				44.8	19.9	18.0
					(47.6)	(34.4)	(25.4)
1	4					27.2	4.5
						(19.9)	(12.5)
p-r-s		5	4	3	2	1	0

Table A2.7: I(2) Trace Test Australia, P2

p-r	r	$S^{\infty}_{r,s}$					$S^{\infty}_{r,p-r}$
5	0	**248.0**	**194.8**	**157.1**	**130.7**	**108.7**	**99.4**
		(198.2)	(167.9)	(142.2)	(119.8)	(101.5)	(87.2)
4	1		**160.7**	**132.2**	**102.4**	**81.2**	**64.1**
			(137.0)	(113.0)	(92.2)	(75.3)	(62.8)
3	2			**105.3**	**76.4**	**54.1**	36.8
				(86.7)	(68.2)	(53.2)	(42.7)
2	3				50.8	22.1	18.2
					(47.6)	(34.4)	(25.4)
1	4					10.6	5.3
						(19.9)	(12.5)
p-r-s		5	4	3	2	1	0

Table A2.8: Long-run Structure Australia

	period 1		*period 2*	
variable	$\widehat{\beta}_1$	$\widehat{\beta}_2$	$\widehat{\beta}_1$	$\widehat{\beta}_2$
Δp	0.198	-1.037	0	0
(s.e.)	(0.020)	(0.106)		
y	-1	0	-1	1
(s.e.)				
$(m2-p)$	1	-3.193	1	0
(s.e.)		(0.196)		
i^s	-0.124	1	0.140	0
(s.e.)	(0.015)		(0.011)	
i^l	+0.124	-1	-0.074	0.006
(s.e.)	(0.015)		(0.011)	(0.002)
trend	0	0	-0.003	-0.005
(s.e.)			(0.000)	(0.000)
	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$
Δp	-0.872	0.655	-13.541	-5.077
(t-rat.)	(-0.380)	(1.455)	(-2.300)	(-0.345)
y	-0.047	-0.010	0.122	-0.694
(t-rat.)	(-1.062)	(-1.147)	(1.640)	(-3.742)
$(m2-p)$	-0.141	-0.032	-0.553	-0.700
(t-rat.)	(-1.943)	(-2.219)	(-2.496)	(-1.263)
i^s	-2.234	-0.485	-4.013	1.253
(t-rat.)	(-1.650)	(-2.824)	(-5.443)	(0.679)
i^l	1.138	0.217	-1.762	16.787
(t-rat.)	(1.728)	(1.677)	(-1.147)	(4.372)
	$\chi^2(3)$=3.64 (0.46)		$\chi^2(3)$=2.41 (0.49)	

Table A2.9: Long-run Structure New Zealand

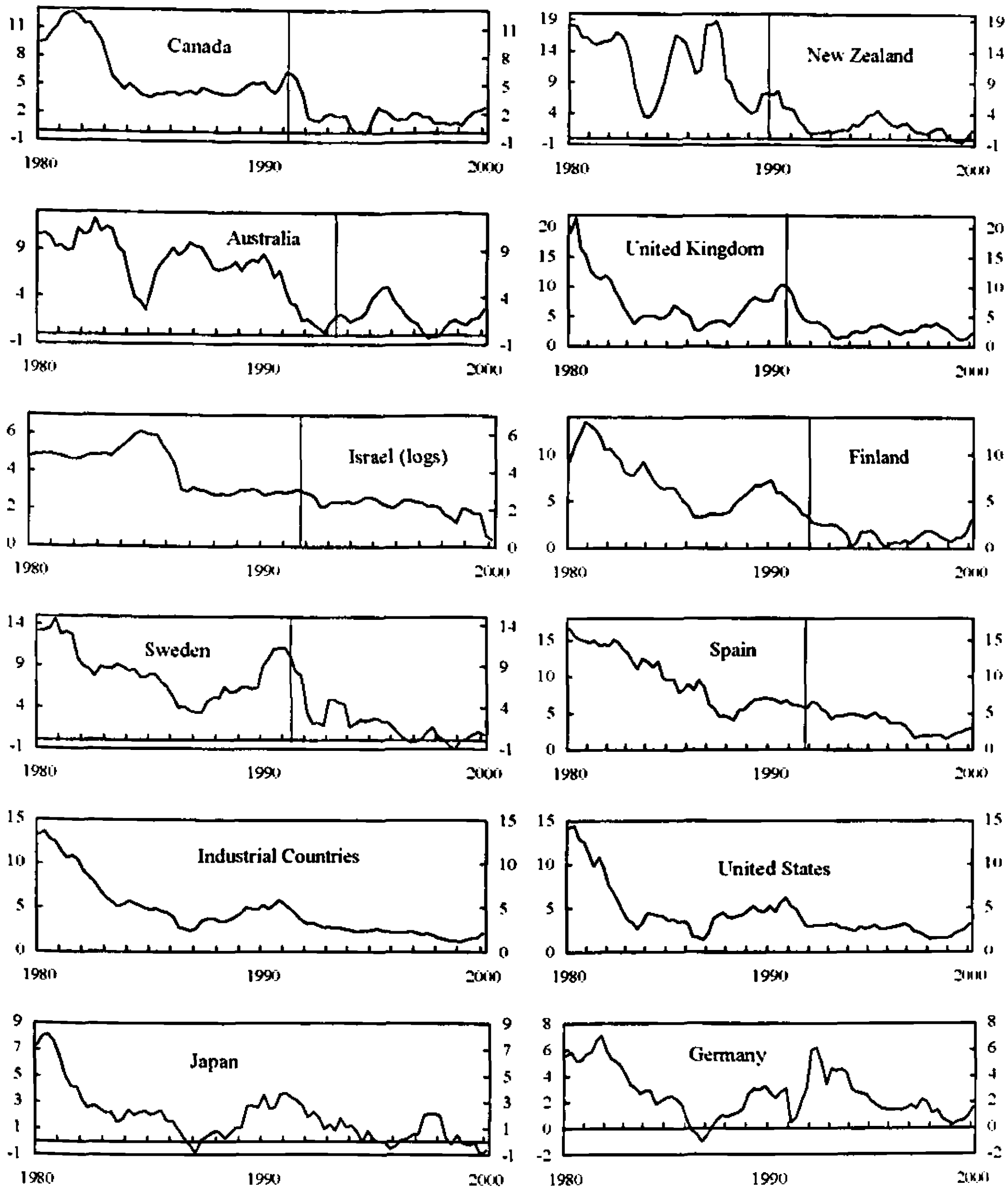
	period 1		*period 2*	
variable	$\widehat{\beta}_1$	$\widehat{\beta}_2$	$\widehat{\beta}_1$	$\widehat{\beta}_2$
Δp	1.160	-0.391	0.078	0
(s.e.)	(0.142)	(0.054)	(0.010)	
y	-1	-4.044	-1	0
(s.e.)		(0.364)		
$(m2-p)$	1	0	1	-3.690
(s.e.)				(0.583)
i^s	-2.153	1	-0.163	1
(s.e.)	(0.080)		(0.008)	
i^l	+2.153	-1	0	0
(s.e.)	(0.080)			
trend	0	0	-0.003	0.028
(s.e.)			(0.000)	(0.004)
	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$
Δp	-2.726	-6.437	-13.457	-2.383
(t-rat.)	(-4.231)	(-3.959)	(-5.288)	(-2.167)
y	0.006	0.015	-0.041	-0.016
(t-rat.)	(0.307)	(0.290)	(-0.334)	(-0.928)
$(m2-p)$	-0.142	-0.325	-0.556	-0.066
(t-rat.)	(-3.287)	(-2.984)	(-3.583)	(-2.820)
i^s	-0.746	-2.762	-3.800	-0.816
(t-rat.)	(-3.004)	(-4.409)	(-1.908)	(-2.698)
i^l	0.084	-0.251	-1.748	-0.346
(t-rat.)	(-0.459)	(-0.541)	(-1.504)	(-1.961)
	$\chi^2(4)$=7.514 (0.13)		$\chi^2(3)$=2.55 (0.47)	

Table A2.10: Long-run Structure Canada

	period 1			period 2	
variable	$\widehat{\beta}_1$	$\widehat{\beta}_2$	$\widehat{\beta}_3$	$\widehat{\beta}_1$	$\widehat{\beta}_2$
Δp	0.566	0	0	1.448	-2.682
(s.e.)	(0.049)			(0.157)	(0.291)
y	-1	0	0	-1	0
(s.e.)					
$(m2-p)$	1	0	0	1	-1.258
(s.e.)					(0.106)
mci	0	-1.049	-0.599	0	0
(s.e.)		(0.179)	(0.151)		
i^s	-0.591	1	0	-0.495	1
(s.e.)	(0.199)			(0.040)	
i^l	+0.591	-1	1	+0.495	-1
(s.e.)	(0.199)			(0.040)	
trend	-0.001	0	0	<0.000	0
(s.e.)	(0.000)			(0.000)	
	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_3$	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$
Δp	-1.807	-1.885	-2.990	2.363	1.650
(t-rat.)	(-8.302)	(-1.999)	(-2.460)	(0.939)	(1.215)
y	-0.002	0.013	0.066	0.085	0.043
(t-rat.)	(-0.353)	(0.516)	(2.109)	(1.648)	(1.538)
$(m2-p)$	-0.004	-0.078	-0.121	-0.550	-0.290
(t-rat.)	(-0.453)	(-2.002)	(-2.430)	(-3.767)	(-3.677)
mci	0.001	-0.025	-0.041	-0.086	-0.050
(t-rat.)	(0.267)	(-1.117)	(-1.423)	(-1.813)	(-1.959)
i^s	-0.010	-0.670	-0.648	0.019	0.005
(t-rat.)	(-0.405)	(-6.151)	(-4.612)	(0.107)	(0.049)
i^l	0.037	-0.004	-0.093	0.221	0.122
(t-rat.)	(2.691)	(-0.073)	(-1.220)	(1.561)	(1.593)
	$\chi^2(7)$=11.84 (0.11)			$\chi^2(5)$=8.74 (0.13)	

2.7.2 Figures

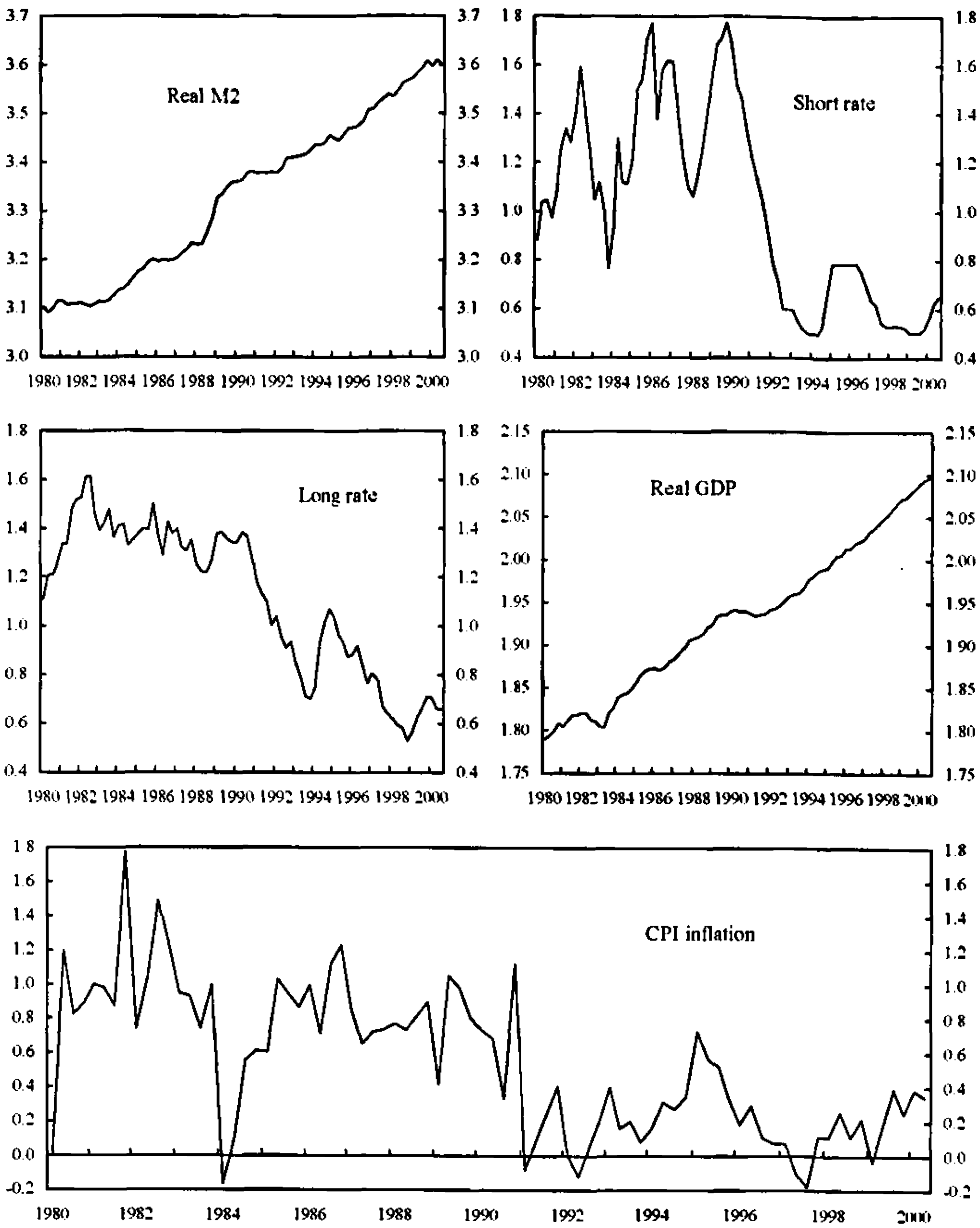
Figure A2.1: 4-Quarter CPI Inflation Rates in Selected IT and Non-IT Countries, 1980-2000



Source: International Monetary Fund, *IFS*.

Notes: Vertical lines signify the approximate introduction of IT.
The chart for Israel is on a logarithmic scale due to a period of hyperinflation.

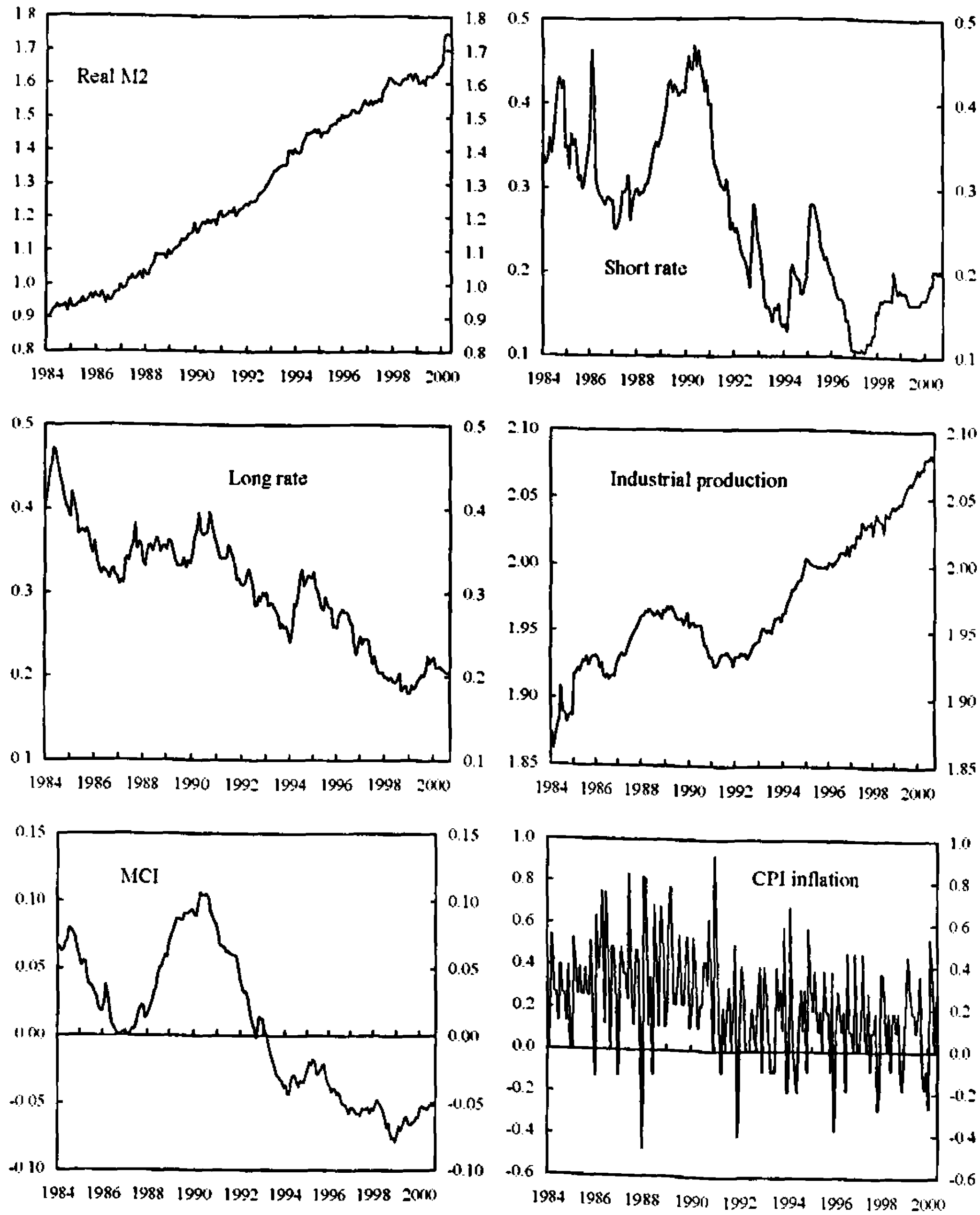
Figure A2.2: Australia: Macroeconomic Data, 1980–2000



Source: International Monetary Fund, *IFS*.

Note: Quarterly data, in (log) levels, CPI in first differences.

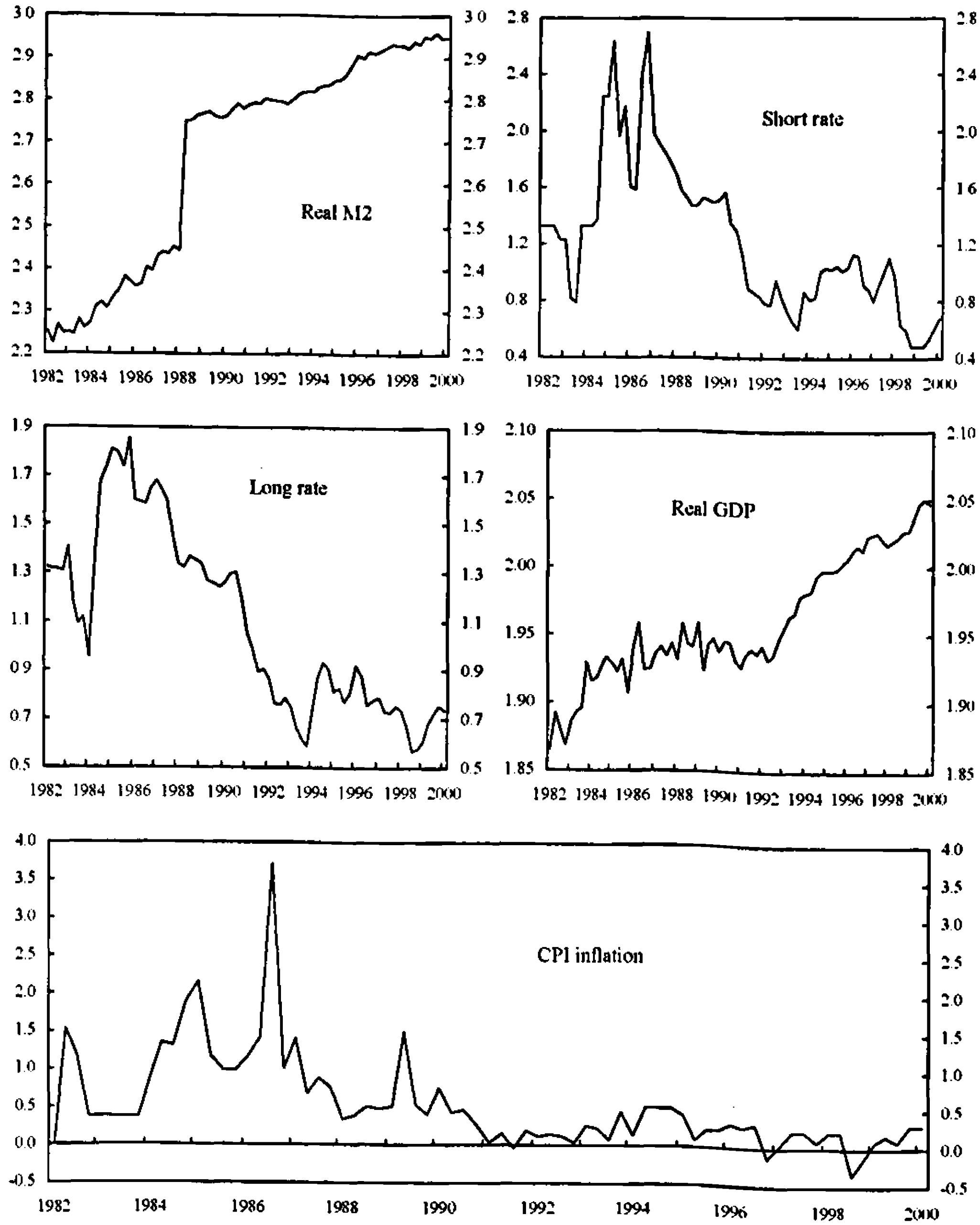
Figure A2.3: Canada: Macroeconomic Data, 1984–2000



Source: International Monetary Fund, *IFS*.

Note: Monthly data, in (log) levels, CPI in first differences.

Figure A2.4: New Zealand: Macroeconomic Data, 1982–2000



Source: International Monetary Fund, *IFS*.

Note: Quarterly data, in (log) levels, CPI in first differences.

Chapter 3

Current Account Fluctuations: How Important Are Nominal Shocks?

3.1 Introduction

Since the publication of the Redux paper (Obstfeld and Rogoff, 1995), the general equilibrium approach to international macroeconomics, often labeled as the "New Open Economy Macroeconomics" (NOEM, for short), has gained remarkable momentum.[1] Roughly speaking, this new paradigm marries microfounded economies open to international exchange with frictions and rigidities stemming from imperfect competition in product or factor markets. This paradigm, in fact, parallels to some extent the closed-economy New Keynesian approach.[2] Up to the present, the literature has mainly taken a theoretical perspective, leaving empirical issues aside. A central, but so far underdeveloped topic in this line of research is the behavior of the current account.

[1]See Vanhoose (2004) and especially Lane (2001a) for valuable surveys.

[2]See, e.g., Clarida *et al.* (1999).

The emphasis of the theoretical NOEM research program has been on two-country models, thus enabling the analysis of international transmission, which occurs mainly through the Keynesian expenditure switching effect. In the Redux model, for example, an unanticipated, expansionary domestic monetary policy shock (the "Dornbusch experiment"[3]) triggers an instantaneous depreciation and a current account surplus. This result is also the outcome of the textbook Mundell-Fleming model of flexible exchange rates, where the depreciation triggered by the monetary expansion redirects international demand towards domestically produced goods. In fact, the basic Mundell-Fleming model does not foresee adverse current account behavior.[4]

The NOEM literature, instead, also provides examples of two-country models, which predict the current account to be in balance (implying long-run money neutrality) or even in deficit in the short run after a monetary shock—see below.

Theoretical small open economy (SOE) models, instead, are still relatively rare in the NOEM literature. Again, Obstfeld and Rogoff (1995) provide the starting point for the analysis. In their model however, the current account is always in equilibrium, even in the short run. Lane (2001b) presents a SOE model that allows for short-run current account imbalances. He offers only limited empirical support for his SOE model, though: the analysis of U.S. data reveals a "J-curve effect," i.e., the current account moves from an initial deficit into a prolonged surplus. However, given that the United States is clearly not a good example of a small open economy, the issue merits further exploration.

This chapter intends to augment the limited empirical evidence on current account behavior in the NOEM literature. Using a vector autoregressive (VAR) framework, the effect of monetary (or better nominal) shocks on the current account will be analyzed for the G-7 countries. To identify nominal shocks, I take advantage of the theoretical prediction of this family of NOEM models, namely that in the long-run,

[3]See Dornbusch (1976); Rogoff (2002) provides an interesting review.

[4]This said, extensions of the original Mundell-Fleming framework have been proposed that allow for adverse effects, such as a J-curve, see Argy (1994), chapter 21.

there is a zero effect, while the short-run reaction is not restricted. In this sense, I do not test a particular prediction for the current account, but take a comparative approach, not necessarily related to model fundamentals: What is the role of nominal shocks in the determination of the current account and how does it differ across (G-7) countries?

To answer this question, I construct three time series for each country, namely domestic output relative to "rest of world" (ROW), the current account relative to domestic output, and domestic prices relative to ROW. Identification is achieved by imposing that orthogonal shocks to the last series have no long-run effects on the two former series; these shocks are consequently labeled as "nominal." In other words, I single out idiosyncratic or "asymmetric" nominal shocks. Then, I simulate impulse response functions (IRFs) for each country, detailing the (short-run) effect of a nominal shock on the current account.

The following results emerge from the analysis: First, short-run current account imbalances after nominal shocks are significant. Second, countries' current accounts are found to react differently to nominal shocks. The current account surplus predicted by classical theory is not robust across countries. While Japan, Italy, and probably the United States reveal a J-curve effect, other countries manifest purely cyclical behavior. Hence, the results obtained by Lane (2001b) cannot be confirmed for a broader sample. Third, while the positive effect on the current account (normalized by its standard deviation) is the highest in Canada and Japan, the relative contribution of a nominal shock in explaining current account variance is maximized in France, the United States, and Italy. There is strong evidence of nominal shocks having short-run real effects, but this evidence is heterogenous across countries. Finally, extending the sample to other (non G-7) OECD countries confirms the conclusion reached for the G-7 economies that there is no consistent reaction of the current account to a nominal shock across countries.

The rest of the chapter is structured as follows: In the next section, I compare two major approaches taken in the literature to current account determination—the

Mundell-Fleming model and the NOEM framework—and review briefly the related empirical literature. Section 3.3 discusses the empirical approach taken in this chapter. Section 3.4 compares empirical estimates of the current account adjustment following a nominal shock for the G-7 countries and proposes a couple of extensions, including to non G-7 economies. Section 3.5 concludes and section 3.6 is an appendix, containing additional tables and charts.

3.2 A Look at the Literature

In this section, first I briefly review the "conventional" Mundell-Fleming approach to current account determination; second, I investigate the various predictions for current account behavior stemming from the NOEM literature; and third, I relate the econometric approach to the existing literature.

3.2.1 The Mundell-Fleming Approach

The classic treatment of a small open economy dates back to (at least) Mundell (1962, 1963).[5] In truly seminal work together with Fleming (1962), the role of capital mobility and the exchange rate regime for the effectiveness of monetary and fiscal policy is clarified, 10 years before the actual adoption of flexible exchange rates around the world after the breakdown of Bretton Woods. The workhorse model of international macro for more than three decades was based on the familiar IS-LM framework, augmented by the balance of payments (external equilibrium).[6] In a nutshell, a unexpected positive domestic monetary shock shifts the LM curve out, leading (in a closed economy) to a rise in income and a (temporary) decrease of the

[5]For an investigation into the origins of the "Fleming-Mundell Model," see Boughton (2003).

[6]See any (macro/international) economics textbook for an exposition, e.g., Burda and Wyplosz (2001), Kenen (1994), or Dornbusch *et al.* (2004). For a particularly detailed treatment, see Argy (1994).

domestic interest rate.[7] In the open economy, two further effects are at work: due to the income expansion, higher imports initially worsen the current account.[8] On the other hand, with high capital mobility, capital outflows will press for a devaluation of the domestic currency to clear the foreign exchange market. This, in turn, will trigger an expenditure-switching effect towards home goods, shift the IS curve, and boost domestic income even further. In the extreme case of perfect capital mobility, no interest rate differentials are sustainable, and monetary policy reaches its highest degree of efficiency under flexible exchange rates. The expenditure switching effect rests on one major hypothesis: price flexibility in the local currency. This assumption is in fact hard to overcome in the traditional Mundell-Fleming setting, since firms' behavior is not microfounded.

3.2.2 The NOEM Approach

Obstfeld and Rogoff (1995, 1996) paved the way for explicit microfoundations in the recent international macro literature.[9] In this literature, the practice has mostly been to characterize the model economy explicitly, while equilibrium dynamics are often modeled as a linearized approximation around the steady state.[10]

In parallel to the closed-economy "New Keynesian Approach," real effects of nominal shocks in these models stem from the fact that prices and/or wages are assumed

[7] The shock considered is supposed to be idiosyncratic or "relative" to the rest of world. A similar monetary shock at home and abroad would not create an interest rate differential.

[8] Note the relationship to the elasticities approach of the current account (or the trade balance if we neglect trade in services and investment income flows for a moment): the Marshall-Lerner(-Robinson) condition for a positive reaction of the trade balance conditional on a devaluation states that the sum of price elasticities of domestic and foreign demands for imports has to be larger then unity (see, e.g., Kenen (1994), p. 352). In fact, short-run elasticities have been found to be significantly lower than in the longer run, contributing hence to an adverse short-run reaction of the trade balance; Krugman and Obstfeld (2003) cite corresponding evidence from Artus and Knight (1984).

[9] First traces of microfoundations date back to Svensson and van Wijnbergen (1989).

[10] Very few papers provide closed form solutions for the dynamics of interest, see, e.g., Obstfeld and Rogoff (2000).

to adjust only slowly. One way to impose price rigidity is by having firms set them one period in advance, such that they are "predetermined." This, in turn, implies that adjustment is completed after one period; that is, persistence effects cannot be captured. More elaborated modeling approaches to nominal rigidity in the literature have used staggered price or wage setting following Calvo (1983) and Taylor (1979).[11] Most work in the NOEM context has built on the analytically more straightforward assumption, namely predetermined prices.

Two-country models in the NOEM literature offer, based on explicitly micro-founded behavior of economic agents, a variety of conclusions as to how the current account reacts following a nominal shock—ranging from a temporary current account surplus to the opposite conclusion. First, in the Redux paper by Obstfeld and Rogoff (1995), an idiosyncratic[12] positive monetary surprise in the home country leads, due to sticky prices, to a rise in domestic consumption and production and lowers the world real interest rate. With nominal depreciation, domestic as well as international demand will (under standard assumptions) switch to domestically-produced tradables, triggering a short-run current account surplus. In this setting, money is not neutral in the long run, since the temporary current account surplus implies an improvement in the net foreign asset position of the home country vis-à-vis the rest of the world. In turn, the net investment inflow allows for a continuous trade balance deficit (domestic consumption higher than production), while the current account is balanced in the new steady state. Furthermore, the wealth effect stemming from the improved net foreign asset position tends to reduce labor supply and, hence, domestic output. This wealth effect is only of second-order importance, however, compared to the positive demand-side effect on output caused by slowly-adjusting prices.

Second, Betts and Devereux (2000) entitle a fraction of firms to discriminate

[11]Closed-economy New Keynesian models that capture persistence effects by making use of staggered price and wage setting have been proposed, e.g., by Galí (2001) and Erceg *et al.* (2000).

[12]Similarly to the Mundell-Fleming framework, an equiproportionate money supply increase would not have any terms-of-trade or exchange rate effects, nor would this type of symmetric shock affect the current account; see Obstfeld and Rogoff (1996), p. 683.

between markets and "price to market" by setting a price in the local currency, at home as well as abroad (local currency pricing, LCP). In this case, the monetary shock still provokes a depreciation of the domestic currency, but no longer triggers price adjustment in accordance with the law of one price, which only holds *ex ante* in this environment. Consequently, the higher the fraction of discriminating firms, the lower the international expenditure switching effect, and the smaller the change in consumption of the domestic traded good since relative prices do not change. In the limit, home and foreign consumption growth are completely delinked, and the current account is unaffected by a monetary policy shock even in the short run. Obstfeld and Rogoff (2000) have criticized this approach; they prefer the assumption of producer currency pricing (PCP). Others have sided with Betts and Devereux (2000) in assuming local (that is, destination market) currency pricing; see, e.g., Devereux and Engel (1998), and Engel (2002).[13]

Finally, Chari *et al.* (2002) obtain yet another prediction for the current account in a two-country model. The introduction of capital as a second factor of production provides an additional effect of the interest rate decline following the monetary expansion. An investment boom will (partly) lead to increased imports, and limit the positive effect of the expenditure switching on the current account. A strong enough investment effect will even lead to a temporary current account deficit.

SOE models, instead, are (so far) rare in the NOEM literature. Obstfeld and Rogoff (1995) also provide a SOE version of their two-country model in the appendix of the Redux paper.[14] They sketch an economy with two sectors, traded vs. non-traded goods. While in the former goods prices obey the law of one price, the non-traded sector is assumed to exhibit nominal (price) rigidities to allow for real effects of monetary policy. Furthermore, the non-traded sector produces a differentiated good under

[13]See also Bergin (2004). Koren *et al.* (2004) find overwhelming empirical evidence in favor of the PCP assumption.

[14]See Bergin (2003) for another axample of a SOE model; he focuses on pass-through and local currency pricing.

monopolistic competition à la Dixit and Stiglitz (1977), while the internationally-traded good is homogenous. The world interest rate is exogenous and the endowment of tradables taken as given by the domestic economy. Hence, the current account is determined over time by the consumption of tradables. The crucial feature of this model is that consumption of tradables and non-tradables enter the utility function log-separably, not allowing consumption of non-traded goods to affect the marginal utility of traded goods' consumption. Therefore, the time path of tradables consumption is flat even after a monetary shock, leading to a current account in perfect equilibrium, even in the short run.

The assumption of log-separability in consumption is relaxed by Lane (2001b). He also distinguishes between a traded, competitive and a non-traded, monopolistically competitive sector. The distinctive feature of this model is a more general formulation of consumption behavior, where consumption of tradables and of non-tradables enter the period utility function in an aggregated way. Possible spillover effects between the non-traded and the traded sector then determine whether a monetary shock affects the current account in the short run and in what direction. Under flexible exchange rates, a current account disequilibrium goes hand in hand with an opposite capital account disequilibrium; that is, net capital imports/exports. In the long run, however, the capital account (and hence the current account) is in equilibrium.[15]

More precisely, a monetary shock in Lane's model triggers a temporary boom in the non-traded sector (since in the short run firms have an incentive to produce more due to the positive mark-up stemming from monopolistic competition, and output is therefore demand-determined). Simultaneously, prices of traded goods rise, in the short as well as in the long run, and the nominal exchange rate depreciates. Whether this short-run boom in the non-traded sector has effects on the traded sector as well as real long-run effects in general hinges on the net outcome of two opposite

[15]This result is not specific to the model proposed by Lane (2001b) but a very general outcome in this literature: A transversality/no-Ponzi-game condition ensures that the long-run capital account is balanced, and therefore (under flexible exchange rates) the current account as well.

substitution effects: On the one hand, traded goods become relatively more expensive due to the monetary shock (the price for non-traded goods is fixed in the short run), and consumers *intra*temporally substitute consumption away from traded goods. On the other hand, the *inter*temporal substitution effect makes consumption today more attractive, since price adjustment in the non-traded sector in the next period leads to a higher steady-state price level.

If the *intra*temporal substitution effect is relatively strong, consumption of tradables is in fact reduced. This, in turn, leads to an immediate current account surplus, and an accumulation of net foreign assets, since production of the traded goods is given and hence unchanged. In the long run, interest revenues on net foreign assets allow higher steady-state consumption of tradables, and hence a permanent net-exports deficit. Money is therefore not neutral in the long run. The current account instead will be balanced, precisely because net imports will be financed from the net foreign-asset revenue.

On the other hand, if the *inter*temporal substitution effect dominates, higher short-run consumption of tradables will cause a temporary current account deficit while in the long run, the economy runs a trade balance surplus, with consumption of traded goods below the exogenous endowment to finance the debt service due to the initial current account deficit.

Finally, if the two substitution effects cancel out, no spillover from the non-traded to the traded sector manifests, the short-run current account balance is unaffected, and the monetary shock does not have (long-run) real effects.

3.2.3 Empirics

Up to now, relatively little empirical work has been done in the NOEM field. From a philosophical point of view, two approaches are possible: First, unconditional moments derived from a calibrated model can be matched with moments of the actual data.[16] Second, conditional moments can be investigated, for example, in a vector autoregressive (VAR) framework.[17] Bergin (2003, 2004) and Chari *et al.* (2002, 2000) propose the former approach, while examples of the latter include Clarida and Galí (1994), Eichenbaum and Evans (1995), and Lane (2001b). The latter family of papers focuses on the effect monetary shocks have on the exchange rate in the presence of sticky prices. Galí (1999) presents evidence for real shocks.

The empirical approach taken in this chapter was developed by Blanchard and Quah (1989). Since then, numerous applications have built on their seminal work. The two major modifications adopted here—triggered by the application to the open economy—are the inclusion of a third variable, the current account, in the estimation process and the particular construction of the "domestic" time series (i.e., relative to the rest of world). In doing so, I follow Clarida and Galí (1994), who investigate the effects of nominal shocks on the real exchange rate—as opposed to the current account. Prasad (1999) also uses a similar multilateral setup to investigate international business cycle propagation. In his contribution, however, he focuses on the effects of output shocks on the trade balance.

Finally, Ahmed and Park (1994) investigate the role of the trade balance for a number of small open economies in a VAR framework. They do not construct relative time series however; instead, they include U.S. output in all VARs as a proxy for the rest of the world. Furthermore, their identification strategy is somewhat different, since they postulate a zero long-run effect of a nominal shock on the trade balance. This is a special case of my assumption, where the net foreign asset position of a given country allows for capital flows such that the trade balance can be in disequilibrium as long as the effect on the current account is zero.

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[16] In the tradition of the computational experiment undertaken by Kydland and Prescott (1982, 1996).

[17] The introduction of the VAR framework into macroeconomics is commonly associated with Sims (1980).

Summing up, there are a number of theoretical papers in the NOEM literature, which conjecture alternative short-run behavior of the current account conditional on a monetary shock. Empirical evidence is limited, however. The next section attempts to fill this gap.

3.3 Empirical Analysis

In this section, I first describe the empirical approach taken to identify nominal shocks in the VAR analysis. The models described in the preceding section share one characteristic feature: while the short-run reaction of the current account in response to a monetary shock depends on various effects, the long-run effect is zero. This theoretical result will be used in what follows as an identifying restriction in the econometric estimation. By focusing on the long run, I do not restrict short-run behavior: both a current account deficit as well as a surplus are possible. Second, I briefly introduce the data and their construction, and, finally, highlight some of the time series properties of the data. The results from the empirical investigation are presented in the next section.

3.3.1 The Empirical Model and Long-Run Restrictions

The models outlined above present some important conclusions for the conditional behavior of the current account, namely that nominal shocks have a zero long-run effect, but that there is no prior as to the short-run response. In this section, I briefly outline how the framework by Blanchard and Quah (1989) can be used in this connection. In accordance with the literature, I focus on time series for output and prices, which are relative to the rest of the world. In doing so, I am able to single out idiosyncratic behavior. In fact, NOEM models do not foresee any real effects of nominal shocks that are common across countries.

I assume that there are three types of (uncorrelated) disturbances that possibly affect the three time series (log) relative output (Y/Y^*), current account (CA/Y), and (log) relative price level (P/P^*), where the * indicates an aggregate variable representing the rest of the world. To identify these disturbances, I make the following assumptions: no disturbance has long-run effects on the time series employed in the estimation, more precisely on the first differences of the original time series (i.e., growth rates are stationary). Furthermore, disturbances to (the growth rate of) relative output might have long-run effects on the level of all three series, while shocks to the (growth rate of the) current account do not have long-run effects on the level of relative output. Finally, disturbances stemming from the (growth rate of) the relative price level, only affect the price level in the long run. These assumptions technically identify the shocks. Given the chosen structure, it seems natural to label the shocks as supply, real demand/absorption, and monetary/nominal, respectively, although the naming convention is irrelevant for the argument, as noted by Blanchard and Quah (1989). The authors also show that small violations of the identification scheme, for example, lasting effects on output stemming from nominal shocks through a "wealth effect" are of minor consequence.[18] In particular, I am interested in the effects of the last shock on the second variable, i.e., a monetary shock on the current account.

This set-up is close to the system studied by Clarida and Galí (1994)—but deviates from it in three respects: (i) as "demand" variable, I focus on the current account rather than the real exchange rate; (ii) the constructed relative series in their paper are bilateral series, that is, they relate the domestic series to only one other country whereas I construct an aggregate rest of the world, consisting of the other G-7 economies;.(iii) in choosing relative prices as my nominal variable —as opposed to, e.g., relative money stocks/growth or the nominal exchange rate as in Clarida and Galí (1994)—I ensure consistency with Lane's (2001b) model in order to reproduce his result for the United States before exploring the implications for a larger sample.[19]

[18]In light of recent U.S. data, potential long-run effects of a growing current account deficit on the level of relative output also come to mind.

[19]Moreover, detrending the series would have highlighted the cyclical effects better. Again, I omit

Let X_t be the stationary vector $(\Delta(Y/Y^*), \Delta(CA/Y), \Delta(P/P^*))'$, and $\varepsilon_t = (\varepsilon_s, \varepsilon_d, \varepsilon_n)'$ the vector of disturbances, where Δ is the difference operator. It follows from the assumptions that X can be written in the following structural representation:[20]

$$X_t = A_0\varepsilon_t + A_1\varepsilon_{t-1} + ... = \sum_{k=0}^{\infty} A_i\varepsilon_{t-k} = A(L)\varepsilon_t, \qquad Var(\varepsilon_t) = I \tag{3.1}$$

where L is the lag operator. The sequence of long-run coefficients can be denoted as $A(1) = A_0 + A_1 + ...$. Element $a_{ij}(k)$ of the matrix A_k is the effect of ε_j on series Δi after k periods, hence the accumulated $\sum_{l=0}^{k} a_{ij}(l)$ gives the effect on the level of series i after k periods. Identification of the shocks is therefore achieved by imposing restrictions of the type

$$\sum_{l=0}^{\infty} a_{ij}(l) = 0 \tag{3.2}$$

for some elements in the 3×3 matrix $A(1)$.

For empirical purposes, the starting point is a reduced-form vector autoregressive (VAR) representation of X_t :

$$B(L)X_t = \eta_t, \qquad var(\eta_t) = \Omega. \tag{3.3}$$

Here, $B(L)$ a 3×3 matrix of lag polynomials, and η_t the reduced-form innovation. Assuming that the roots of the characteristic equation lie outside the unit circle, this expression can be inverted and rewritten in the following moving average (MA) form:

$$X_t = C(L)\eta_t = C_0\eta_t + C_1\eta_{t-1} + C_2\eta_{t-2} + ... \tag{3.4}$$

where $C(L) = B(L)^{-1}$ and $C_0 = I$. Comparing equation (3.1) with equation (3.4),

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this step to stay consistent with Lane (2001b).

[20] In what follows, I omit the constant.

it follows that $\eta_t = A_0\varepsilon_t$, as well as $A_i = C_i A_0\ \forall i$. A convenient way to identify A_0 is by imposing that $A(1)$ be lower triangular.[21] This is where the assumptions about the structural shocks made above come in: the restrictions in equation (3.5) precisely satisfy this structure.

$$A(1) = \begin{pmatrix} \sum_l^{\infty} a_{11}(l) & 0 & 0 \\ \sum_l^{\infty} a_{21}(l) & \sum_l^{\infty} a_{22}(l) & 0 \\ \sum_l^{\infty} a_{31}(l) & \sum_l^{\infty} a_{32}(l) & \sum_l^{\infty} a_{33}(l) \end{pmatrix}. \tag{3.5}$$

Neither real demand (element 1,2) nor nominal shocks (element 1,3) have long-run effects on the accumulated level of output, and nominal shocks have zero long-run consequences for the current account (element 2,3).

3.3.2 The Data

The data I use are quarterly and stem from the IMF's *International Financial Statistics* database. In the first step, the sample comprises the G-7 countries. This choice is driven by the trade-off between data availability and model congruency. Including more countries would stress the underlying small open economy assumption, but, on the other hand, make the construction of the relative series much more cumbersome. Due to non-availability of consistent Japanese current account data, the observation period starts only four years after the breakdown of Bretton Woods and spans from 1977Q1 to 1998Q4, for a total of 88 observations. The last observation is chosen to coincide with the introduction of the euro in three of the G-7 countries, acknowledging the change in policy regime, and providing better comparability with results in the literature.

The times series employed in the estimation are:

- (log) seasonally adjusted domestic output volume (1995=100, IFS line 99BVRZF), relative to ROW;

[21] See Blanchard and Quah (1989) or Clarida and Galí (1994) for proofs of this statement.

- the domestic current account balance (78ALDZF) relative to domestic GDP[22] (line 99B.CZF); and

- the (log) domestic consumer price level (1995=100, line 64...ZF) relative to the price level in the ROW.

For each G-7 economy, the ROW is proxied by the other six G-7 countries and is, hence, country-specific. The corresponding ROW series are constructed as a weighted averages of national series. The weights, in turn, are given by average annual GDP in international prices over the period 1973-88, taken from the Penn World Tables, mark 5.6.[23] The raw current account series has been passed through a X11-filter to remove remaining seasonal cycles.[24]

3.3.3 Estimation Issues

In order for the above setup to be valid, the data series can be non-stationary in levels, but must be stationary as employed in the VAR; that is, in first differences. With very few exceptions, conventional unit root tests have confirmed the visual impression, indicating non-stationarity in levels but not in first differences (see Table A3.2 in the appendix to this chapter for detailed results).

The estimation in first differences ignores possible cointegration between the time series in levels. To check for this possibility, simple trace tests following Johansen (1996) have been implemented. Only Japan and Germany offer signs of cointegration

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[22]The series were converted from USD (as provided in the IFS) into domestic currency by the period average exchange rate over the quarter, taken from the IFS as well (line ..AE.ZF).

[23]See Summers and Heston (1991) for a discussion of mark 5. Table A3.1 in the appendix to this chapter presents the relative weights employed. In constructing the weights, I use annual observations until 1988 to take into account East Germany, for which no observations are available after 1988.

[24]Note that for the construction of relative output, seasonally adjusted series have been employed where possible. In the extended—i.e., non-G7—sample analyzed further below, some countries do not provide seasonally adjusted data, and the X-11 filter has been used to smooth series (lines 99BVP..).

(see Table A3.3 in the appendix).[25] Being aware of the problems deriving from the omission of the cointegrating relationship as a RHS variable, a VAR representation in first differences is chosen for the sake of a consistent specification across countries.

In the same vein, The VARs, which were estimated for each country separately, included eight lags to obtain a specification free of residual autocorrelation and similar across countries.

3.4 Results

In this section, I present the results of the estimation. First, (accumulated) impulse responses of the current account to nominal shocks are compared across countries. Second, the relative importance of supply, demand, and nominal shocks for the current account is described with the help of variance decompositions. Third, I extend the results to a larger sample by including most OECD economies.

3.4.1 Impulse Responses

To construct the impulse responses, every G-7 country is, in turn, assumed to be the "home" country; that is, the rest-of-world aggregates stand for the other six large economies. Figures A3.1–A3.7 in the appendix present the accumulated impulse responses of the current account to a positive one-standard-deviation nominal shock.[26] There, the middle line represents the accumulated response, while the upper and lower line indicate two standard error bands, calculated by Monte Carlo simulations (500 replications).

The IRFs yield a few interesting results. First, all responses are strongly sig-

[25]A rudimentary cointegration analysis offered the following results. In Japan, the hypothesis of the current account being cointegrated with the relative price level could not be rejected at conventional levels. Furthermore, the price level seems to be weakly exogenous. In Germany, cointegration between the three series (and exogeneity of relative output) could not be rejected at conventional levels.

[26]Note the different scale of the graphs.

nificant. The particularly tight shape of the error bands is due to the "terminal condition," namely a zero response (and hence zero error) of the current account in the long run.

Second, the short-run consequences of a nominal shock on the current account are strikingly dissimilar across countries. While Japan, the United States, and Italy display a lasting surplus, other countries—such as Canada and the United Kingdom—exhibit mere fluctuations around a balanced current account. On impact, the (accumulated) impulse response in France and the United Kingdom is positive, negative for the rest. The latter reaction (if leading to a surplus) could be interpreted as a J-curve effect. In other words, an initial deficit after a devaluation is due to the immediate price effect, adjusting the value of imports—while the quantity effect with expenditure switching in favor of domestic goods manifests only sluggishly. Consistent with results in the literature—but contrary to what one would expect—the United States display a pronounced J-curve effect. The current account moves into surplus after roughly four quarters and remains positive (with the exception of one tick) for roughly four years.[27] A similar surplus occurs also in Japan, but without initial deficit. Germany as well moves almost immediately into a quite persistent surplus, lasting for roughly ten quarters. The current account imbalances in Italy (and, to a lesser extent, in the United Kingdom) die out surprisingly slowly, indicating probably the lasting effect of the ERM exit in 1992.

To facilitate comparison between graphs, Table 3.1 presents key statistics.[28] Absolute effects on the current account differ widely. Relative effects instead—that is,

[27] Lane (2001b) finds results for the United States, which are very similar in shape, and only slightly more pronounced in size. This difference is due to somewhat different weights in constructing the ROW series.

[28] Note: "CA_{max}" gives the maximum current account surplus in response to a one-standard-deviation nominal shock (multiplied by 100, i.e., in percent of the ratio current account to GDP). The second and third line put this number into context by relating it to the standard deviation of the country's current account measure. "Duration CAS" gives the duration (in quarters) of the current account surplus in which the maximum effect occurs. The last line gives the maximum duration of a current account surplus, bold indicating the cycle coincides with the maximum surplus. The second longest surplus duration is added in parenthesis.

Table 3.1: Comparing Nominal Shocks

	Country						
	Canada	France	Germany	Italy	Japan	UK	US
$CA_{max}(\times 100)$	0.0419	0.0834	0.1445	0.1319	0.0345	0.0934	0.0107
$\sigma_{CA}(\times 100)$	0.4019	1.2313	2.1325	1.8917	0.3373	1.9344	0.2741
Rel. effect (%)	10.44	6.77	6.78	6.98	10.22	4.83	3.90
Duration CAS	4	8	9	∞	∞	(10)	28
Max. dur. (2nd)	5 (**4**)	12 (9)	**9** (4)			11 (**10**)	**28** (3)

scaled by the standard deviation of the current account measure—are of the same order of magnitude for all seven economies. The strongest responses of the current account occur in Canada and Japan, while figures for continental European countries are surprisingly similar to each other.

In terms of the model presented by Lane (2001b), for most countries no clear interpretation can be given since the inter- and intratemporal elasticities of substitution are not meant to be varying over time; in other words, the current account is predicted to move only in one direction. The striking exemptions are Japan, and to a lesser extent Italy and the United States, where the current account surplus over the full time horizon documents a dominating intratemporal substitution effect. The spillover between non-traded and traded goods is negative, a consumption boom in the non-traded sector does not trigger higher imports. Hence, goods substitution due to the devaluation can play in favor of domestic goods, and a current account surplus is achieved.

Summing up, the dissimilarity of the current account adjustment across countries points to models that do not restrict its behavior (in one direction). It seems that different channels of monetary transmission across countries lead to varying degrees of expenditure switching. Three possible explanations are suggested.

First, some of the G-7 countries analyzed cannot be considered "small." A sustained current account imbalance might have wealth effects on labor supply (and hence relative output) triggered by its absolute size. For example, the U.S. current account deficit in 2000 amounted to US$ 444.7 billion, or 4.5 percent of U.S. GDP, absorbing some 20 percent of worldwide saving. In an extension (section 3.4.3.2), the heterogeneity result will be "tested" in an augmented sample (14 additional OECD countries).

Second, the extent to which goods are priced in local currency bears heavily on the expenditure switching effect and, consequently, on the current account. More pronounced local currency pricing (in the partner country) limits the positive effect of a devaluation on the current account (of the home country). In terms of the above empirical results, exporters from United Kingdom, France, and to some extent Canada appear somewhat more likely to price in local (foreign) currency than the exporters from the substantially stronger export performers United States, Japan, and Germany. In terms of the LCP vs. PCP discussion, this seems a reasonable conclusion that can be explained by a combination of risk aversion on the exporters' side and market power. For example, Donnenfeld and Haug (2003) also find evidence that "size matters;" that is, the larger the exporter's country is relative to the importer's country, the higher is the fraction of transactions invoiced in the former national currency.

Finally, the procedure implicitly assumes that the equilibrium current account is balanced (i.e., zero). It has been shown that this assumption does not necessarily hold true for a large number of economies, most notably the United States, which are prone to run a (sizeable) long-run equilibrium deficit.[29]

3.4.2 Variance Decompositions

In this section, I disentangle the relative importance of supply, absorption, and nominal shocks in explaining the variance of the current account variable. More precisely, the outcomes of the variance decomposition represent the share of variance of the n-step forecast error of a variable that can be explained by the innovation in another

[29] A significant amount of research along these lines has been carried out at the International Monetary Fund, see Isard and Faruqee (1998) and Isard *et al.* (2001), as well as Dunaway *et al.* (2001) for a discussion of the U.S. case. EFN (2002) provides a brief survey.

variable. This decomposition is based on the series used in the estimation process; that is, on the series in *first differences*.[30] Table 3.2 gives a summary of the long-run decomposition (i.e., after 50 quarters) of the current account variable.

Table 3.2: Summary of Current Account Variance Decompositions (Diffs)

Country	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
Canada	0.221	13.23	70.55	16.22
France	0.643	24.39	45.63	29.96
Germany	0.880	34.57	50.85	14.56
Italy	0.925	16.69	60.29	23.00
Japan	0.093	21.70	73.04	5.25
UK	0.957	15.33	76.37	8.29
US	0.079	33.46	42.71	23.82

Across all countries, shocks to the current account itself (labeled "absorption") are the most important in explaining the forecast error, ranging from 76 percent in the United Kingdom (after 50 quarters), to 42 percent in the United States. On the other hand, the importance of nominal shocks for the CA varies from 5 percent (Japan) to 30 percent (France). Over time—that is once the interaction between the variables is felt—the contribution of the nominal shock tends to rise, while the "own effect" of the current account decreases in all countries. This is particularly the case for France, Canada, and Italy. This evidence leads to conclude that shocks identified and labeled as "nominal" can explain, at least in some countries and to a non-negligible degree, the unexpected variance of the current account.[31]

[30] See Tables A3.4–A3.10 in the appendix of this chapter.

[31] I am, however, not able to reproduce the result by Lane (2001b), who found that monetary shocks contributed almost 50 percent to the variation of the U.S. current account at a horizon of up to 20 quarters.

Table 3.3: Summary of Current Account Variance Decomposition (Levels)

Country	Step	Std error (×100)	"Supply"	"Absorption"	"Nominal"
Canada	1	0.158	30.07	59.31	10.62
	50	0.399	37.62	45.69	16.69
France	1	0.483	28.04	45.88	26.08
	50	1.371	26.77	66.61	6.63
Germany	1	0.724	57.28	41.82	0.90
	50	2.201	43.90	44.94	11.16
Italy	1	0.721	6.98	92.62	0.40
	50	1.784	18.53	61.42	20.05
Japan	1	0.0743	0.06	53.03	46.91
	50	0.229	38.70	54.15	7.15
UK	1	0.789	8.65	91.33	0.02
	50	2.047	7.20	91.25	1.55
US	1	0.0635	0.21	27.72	72.07
	50	0.231	30.92	42.90	26.18

3.4.3 Robustness Checks and Extensions

3.4.3.1 Stationarity of the Current Account

The empirical approach used in this chapter is subject to criticism. For example, considering the current account a non-stationary variable is an economically questionable assumption—even though descriptive statistics indicate this property of the time series. On theoretical grounds, a convincing case can be made for *stationarity* of the current account. Employment of first differences, and therefore "overdifferentiation" implies a loss of statistical information. Since in the estimation the first difference has been used, the variance decomposition presented above allocates the forecast variance of the *first difference* of the current account to the various shocks. For this reason, I reestimated the systems employing the current account in levels, which yields the variance decompositions for the *level* of the current account (Table 3.3).

A comparison of Table 3.3 with Table 3.2 in the preceding section reveals that for most countries, the contribution of nominal/monetary shocks in explaining the forecast error of the current account *level* is similar to the contribution in *first differences*.

The striking exception is France: if calculated for the level, monetary shocks are only able to explain about 7 percent of the forecast error at the 50-quarter horizon (instead of 30 percent in differences). This result can be interpreted as indirect evidence of the non-stationary of the French current account series: since the forecast error goes to infinity with the horizon when forecasting a non-stationary series, the contribution of nominal shocks (which are assumed to have no permanent effect) has to go to zero with the forecast horizon. On the other hand, there is no strong *a priori* reason to believe that the French current account is characterized by weaker mean-reversion than the other countries' series.

3.4.3.2 Beyond G-7: True Small Open Economies

In this section, the criticism of the "small-open-economy assumption" is explored more seriously. In fact, the empirical estimation procedure does not rely on this assumption, but since many international economists are still used to thinking in the IS-LM framework and compare effective outcomes to those derived in this context, a word of caution is warranted. The United States is not a good example of a small open economy, nor are Japan and Germany. Therefore, I extend in this section the analysis to other OECD (but not G-7) countries.[32] To keep the data gathering tractable, I make one important assumption: the ROW series in this section are no longer constructed as the preceding section (i.e., excluding only the country in question); instead, the G-7 economies together are taken to represent the rest of world for all smaller countries.[33] This pragmatic approach is motivated and validated

[32]Note that for data availability reasons, we exclude four recent (out of 30) OECD member countries (Czech Republic, Hungary, Poland, and Slovak Republic), as well as Iceland. Furthermore, Luxembourg was dropped because no separate trade data is available due to the currency union with Belgium. The construction of a joint (i.e. Belgium and Luxembourg) series for prices and output provided little additional insight, given the very moderate weight of the latter in term of GDP in international prices (around 4 percent). Greece was eliminated due to missing comparable nominal quarterly GDP data (line 99B) on the IFS tape. The same holds for Ireland, where IFS data starts only in 1997. Norway is dismissed because it lacks current account data for 1992Q1-1993Q4. Data for Switzerland stem partly from the Swiss National Bank. This leaves 14 OECD but non-G7 countries.

[33]The according weights (in percent) are: Canada 4.37, US 45.93, Japan 15.62, France 8.15, Germany 11.19, Italy 7.18, UK 7.56.

by the observation that the G-7 countries account on average for almost 80 percent of the total OECD output in international prices between 1973 and 1988.

In the appendix, Figures A3.8–A3.10 present the current account impulse responses of these 14 additional countries. The message from the empirical assessment is clear: there is no uniform pattern, in particular the J-curve effect found by Lane (2001b) for the United States, and confirmed in the previous section, arises in only two countries, Mexico and the Netherlands. Other countries, instead, exhibit sustained current account deficits after an initial surplus (e.g., Austria, Finland, or Korea), or even no significant surplus at all, as in the cases of Australia, Belgium, Portugal, or Switzerland. Hence no clear pattern emerges with regard to the pricing-to-market discussion in the broader sample.

3.5 Conclusion

Do nominal shocks have real effects, for example on the current account? If if so, are they similar to what Mundell (1963) predicted? To answer this question, the chapter investigated empirically whether relative nominal shocks have a short-run effect on the current account, imposing a long-run neutrality restriction à la Blanchard and Quah (1989) to identify structural shocks.

In the first part of this chapter, I reviewed theoretical contributions to the current account literature, including the fundamentally Keynesian expenditure switching effect in the classic Mundell-Fleming paradigm and the NOEM literature which enjoys growing support.

The empirical investigation in the second part builds on constructed time series that represent the economy's state relative to the rest of the (G-7) world. Due to the long-run identifying restrictions, "nominal shocks" can be separated from "supply" and "real demand/absorption" shocks. I find that the reaction of the current account is profoundly different across countries, ranging from a J-curve effect (United States, Japan, Italy) to purely cyclical movements (United Kingdom, Canada). Furthermore, the importance of nominal shocks in explaining the current account varies across countries. However, it is never the main explanatory component. Extending the sample to other small OECD countries confirms the heterogeneity result.

From this exercise, the need emerges for microfounded (SOE) current account models that stress country-specific particularities. One direction of future research would lead to a more elaborated modelling of nominal rigidities in general. Especially integrating the wage-setting process, which has been found to be particularly heterogenous across countries, could shed further light on the question why current account reactions are so different across countries. A second line of research could investigate further the somewhat inconclusive evidence regarding the interaction of producer vs. local currency pricing and the current account by integrating the small open economy framework better with the pricing-to-market paradigm to yield empirically testable implications, for example along the lines of Bergin (2004), Choudri *et al.* (2002), and Monacelli (1999, 2003).

3.6 Appendices

3.6.1 Tables

Table A3.1: GDP Weights for ROW Countries

	Relative weight of G-7 partner country						
Home country	Canada	US	Japan	France	Germany	Italy	UK
Canada		0.480	0.163	0.085	0.117	0.075	0.079
US	0.073		0.274	0.158	0.216	0.129	0.151
Japan	0.052	0.544		0.097	0.133	0.085	0.090
France	0.048	0.500	0.170		0.122	0.078	0.082
Germany	0.049	0.517	0.176	0.092		0.081	0.085
Italy	0.047	0.495	0.168	0.088	0.121		0.081
UK	0.047	0.497	0.169	0.088	0.121	0.078	

Note: Weights are constructed using data on per-capita GDP in international prices and population figures for the period 1973-88, taken from the Penn World Tables, mark 5.6. Germany includes figures for East and West.

Table A3.2: Time Series Properties of the Data (Stationarity)

	Country						
	Canada	France	Germany	Italy	Japan	UK	US
Levels							
$log\,(Y/Y^*)$	-1.95	-2.78	-2.13	-3.97*	-2.07	-3.51*	-1.53
(CA/Y)	-2.22	-3.10	-1.79	-3.37	-2.63	-1.76	-1.76
$log(P/P^*)$	-0.74	-1.99	-2.69	-3.12	-2.47	-2.80	-3.96*
Differences							
$dlog\,(Y/Y^*)$	-3.16*	-3.96**	-2.92*	-3.75**	-4.35**	-3.37*	-3.62**
$d(CA/Y)$	-5.99**	-4.16**	-3.53**	-3.93**	-4.27**	-4.31**	-3.44*
$dlog(P/P^*)$	-3.08*	-2.94*	-2.87	-2.67	-3.69**	-2.97*	-3.27*

Note: The series were tested for a unit root using augmented Dickey-Fuller tests, implemented in PcGive. The null hypothesis is that of a unit root. In the error correction specification, this is rejected if the coefficient on the lagged variable is negative and significantly different from zero. In the table, the test statistics of the coefficients are displayed. 1 or 2 stars exhibit significance at the 5 or 1 percent level. The critical values do not follow the conventional t-distribution. For the estimation in levels, where a constant and a trend have been included, the critical values are -3.47 (5 percent) and -4.08 (1 percent). For the unit root test of first differences, only a constant has been included, and the corresponding critical values are -2.90 (5 percent) and -3.52 (1 percent).

Table A3.3: Cointegration Properties of the Series in Levels

H_0	Country							Trace95
	Canada	France	Germany	Italy	Japan	UK	US	
$r = 0$	25.25	22.34	**39.78**	28.87	**35.93**	20.75	21.78	29.4
$r \leq 1$	13.74	9.71	10.46	11.68	11.41	6.14	5.16	15.3
$r \leq 2$	4.44	3.07	4.51	2.06	0.22	1.75	0.07	3.8
Eigenvalues								
λ_1	0.148	0.160	0.334	0.212	0.288	0.183	0.206	
λ_2	0.121	0.088	0.079	0.125	0.144	0.059	0.068	
λ_3	0.059	0.041	0.060	0.028	0.003	0.024	0.001	

Note: The trace statistic seeks to identify how many eigenvalues are significantly different from zero, in other words the number of dimensions of the cointegrating space. $H_0(r)$ indicates the null hypothesis, the alternative being $H_1(r+1)$. Here, λ_i gives the eigenvalues obtained in the process of the maximization of the likelihood function. They correspond to squared canonical correlation coefficients and indicate, roughly speaking, the degree of correlation between the stationary part of the system and the potentially stationary cointegrating vector. λ_1^{CN}=0.148 indicates therefore a correlation of approximately 38 percent. Values in the tables indicate the test statistic from the PcFIML output (rejections in bold), and Trace95 gives the critical value, more precisely the 95-percent quantile of the likelihood ratio test for the cointegrating rank, taken from Johansen (1996), Table 15.3.

Table A3.4: Variance Decomposition of the Current Account (Canada)

Step	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.165	0.71	96.19	3.10
2	0.170	5.17	91.73	3.10
3	0.172	6.10	89.32	4.58
4	0.177	6.17	89.51	4.31
5	0.187	7.85	87.02	5.13
6	0.196	7.88	81.42	10.70
7	0.202	12.47	76.83	10.70
8	0.202	12.44	76.92	10.65
9	0.205	13.67	75.32	11.01
10	0.211	13.27	74.53	12.2
⋮	⋮	⋮	⋮	⋮
50	0.221	13.23	70.55	16.22

Note: In Table A3.4-A3.10, the variance decomposition of the G-7 current accounts are presented. To preserve space, the tables have been truncated once reasonable convergence to the long-run value was reached. For further elaboration, please see text, section 3.4.

Table A3.5: Variance Decomposition of the Current Account (France)

Step	Std.error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.484	18.14	69.44	12.42
2	0.563	18.55	53.95	27.48
3	0.570	18.08	53.93	27.98
4	0.584	21.01	51.90	27.08
5	0.592	20.96	51.08	27.94
6	0.607	23.26	48.56	28.17
7	0.621	23.68	47.55	28.76
8	0.624	23.58	47.87	28.53
9	0.626	23.80	47.77	28.43
10	0.628	23.59	47.37	29.04
⋮	⋮	⋮	⋮	⋮
50	0.643	24.39	45.63	29.96

Table A3.6: Variance Decomposition of the Current Account (Germany)

Step	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.654	25.75	69.47	4.77
2	0.693	25.80	63.51	10.68
3	0.697	26.21	62.88	10.90
4	0.705	25.90	62.30	11.79
5	0.708	26.19	61.85	11.95
6	0.742	32.40	56.31	11.28
7	0.760	32.18	53.70	14.11
8	0.770	33.91	52.33	13.75
9	0.812	31.16	55.42	13.41
10	0.818	31.28	54.51	14.21
⋮	⋮	⋮	⋮	⋮
50	0.880	34.57	50.85	14.56

Table A3.7: Variance Decomposition of the Current Account (Italy)

Step	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.722	6.17	76.20	17.62
2	0.789	8.29	67.30	24.40
3	0.798	9.86	66.08	24.04
4	0.831	16.26	61.04	22.69
5	0.832	16.35	60.97	22.67
6	0.834	16.35	60.82	22.82
7	0.835	16.30	60.81	22.87
8	0.841	16.40	60.72	22.87
9	0.869	15.43	62.89	21.67
10	0.887	14.80	62.31	22.88
⋮	⋮	⋮	⋮	⋮
50	0.925	16.69	60.29	23.00

Table A3.8: Variance Decomposition of the Current Account (Japan)

Step	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.0784	6.17	93.74	0.08
2	0.0835	12.39	85.84	1.76
3	0.0863	15.19	80.49	4.31
4	0.0869	16.13	79.32	4.54
5	0.0875	15.93	79.53	4.53
6	0.0877	15.85	79.53	4.61
7	0.0892	17.86	77.44	4.69
8	0.0908	20.03	74.90	5.06
9	0.0912	19.86	75.10	5.02
10	0.0919	20.43	74.51	5.05
⋮	⋮	⋮	⋮	⋮
50	0.0934	21.70	73.04	5.25

Table A3.9: Variance Decomposition of the Current Account (UK)

Step	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.794	8.39	90.88	0.71
2	0.826	9.00	90.16	0.83
3	0.827	9.07	90.08	0.83
4	0.843	9.23	86.93	3.83
5	0.873	13.65	81.18	5.16
6	0.875	13.71	81.15	5.13
7	0.886	14.61	79.83	5.54
8	0.893	14.47	79.76	5.75
9	0.926	14.41	78.98	6.59
10	0.932	14.31	79.17	6.51
⋮	⋮	⋮	⋮	⋮
50	0.957	15.33	76.37	8.29

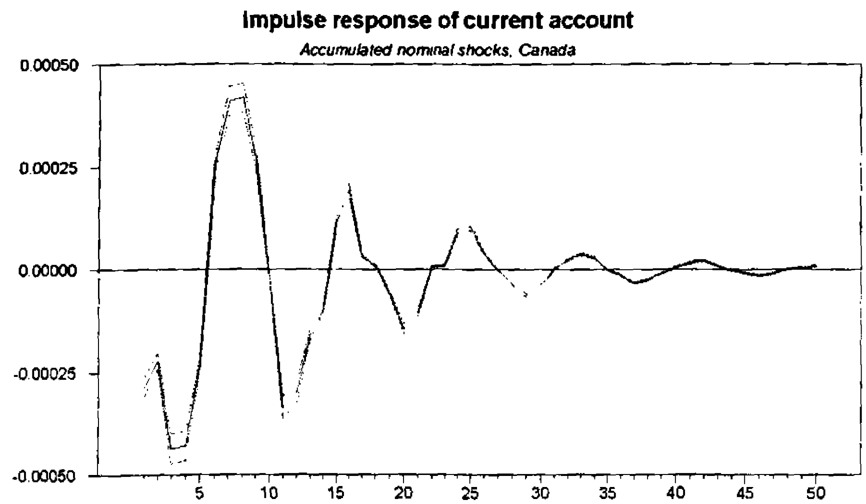
Table A3.10: Variance Decomposition of the Current Account (US)

Step	Std. error (×100)	"Supply"	"Absorption"	"Nominal"
1	0.0668	33.72	45.06	21.21
2	0.0691	34.33	45.52	20.14
3	0.0707	34.18	45.65	20.16
4	0.0738	34.44	43.16	22.40
5	0.0747	33.82	43.47	22.71
6	0.0752	33.65	43.09	23.26
7	0.0772	34.13	43.81	22.06
8	0.0780	33.59	43.25	23.15
9	0.0782	33.42	43.33	23.24
10	0.0787	33.48	42.93	23.59
⋮	⋮	⋮	⋮	⋮
50	0.0791	33.46	42.71	23.82

3.6.2 Figures

3.6.2.1 G-7 Countries

Figure A3.1: Accumulated Impulse Response of the Canadian Current Account to a Nominal Shock



Note: In Figures A3.1–A3.7, the middle line represents the (accumulated) response of the current account to a one standard deviation nominal shock as identified by the long-run restrictions. The two outer lines are ± 2 standard error bands, calculated by Monte Carlo simulations (500 replications).

Figure A3.2: France

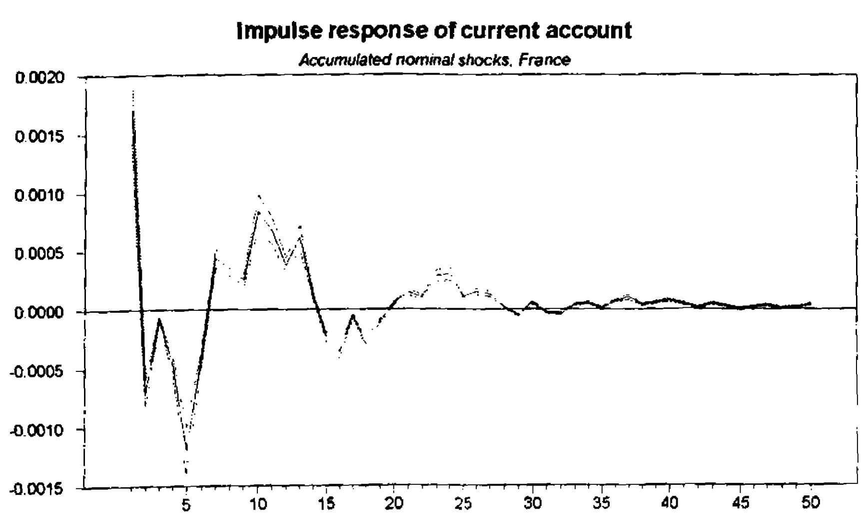


Figure A3.3: Germany

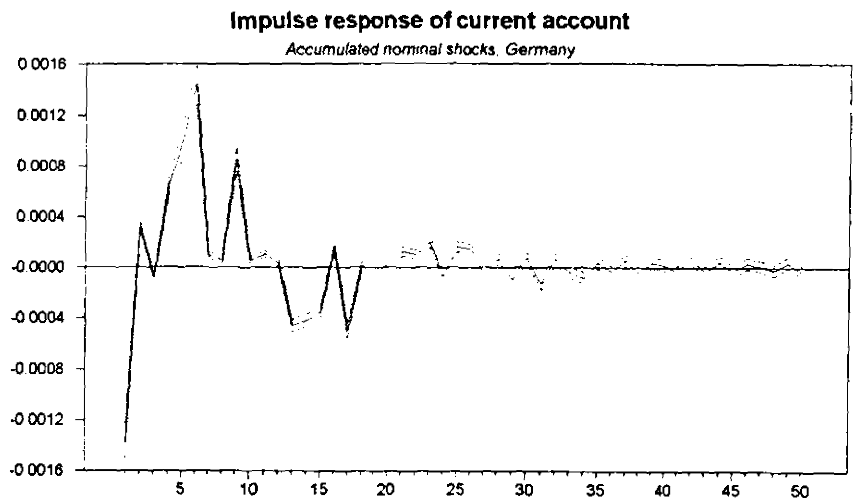


Figure A3.4: Italy

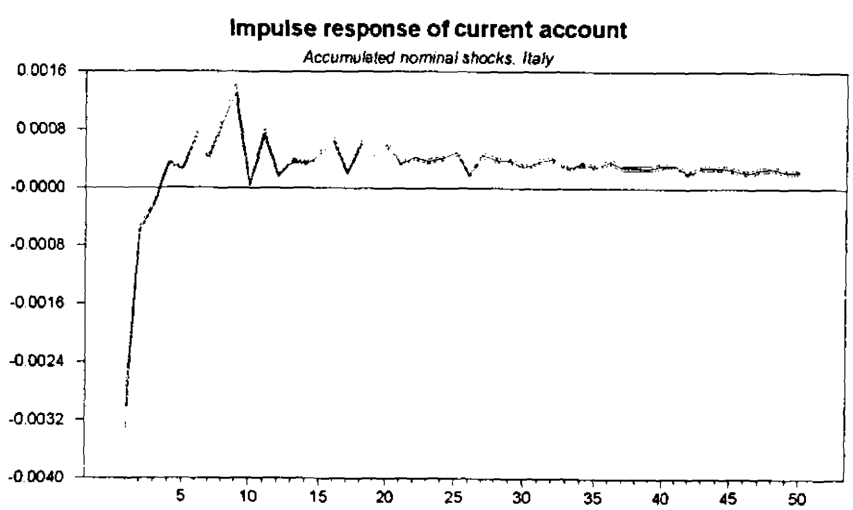


Figure A3.5: Japan

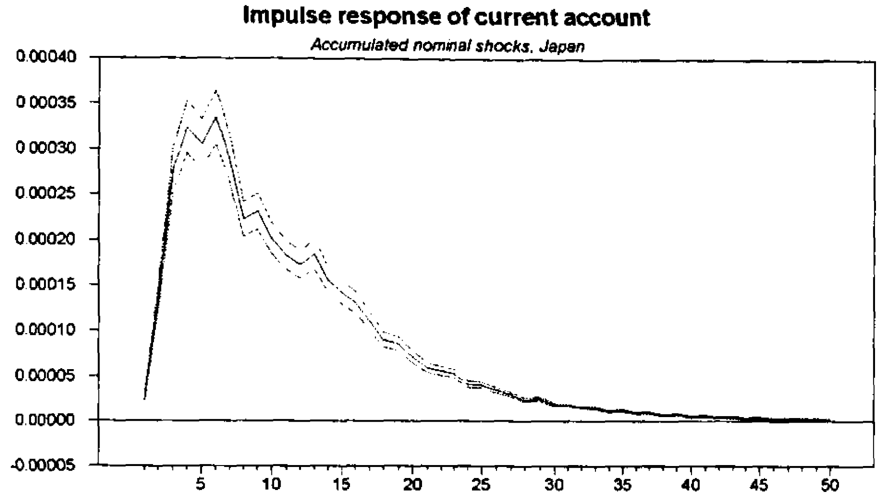


Figure A3.6: United Kingdom

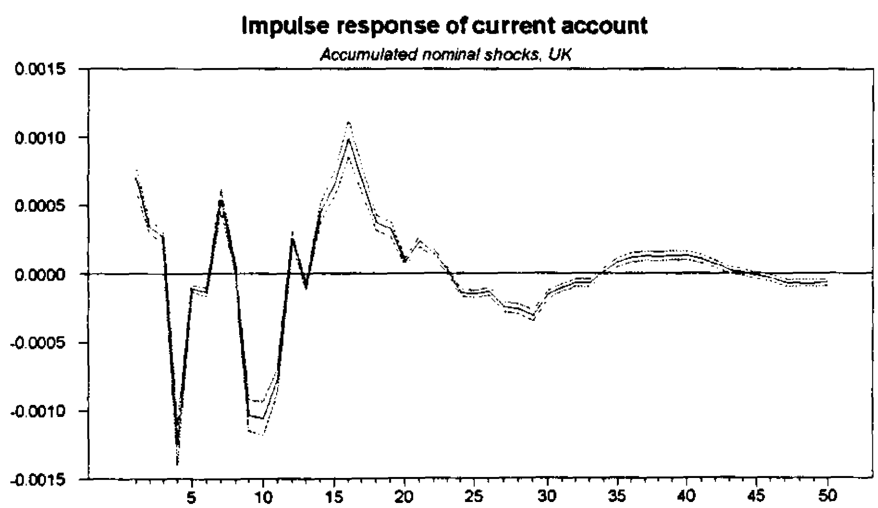
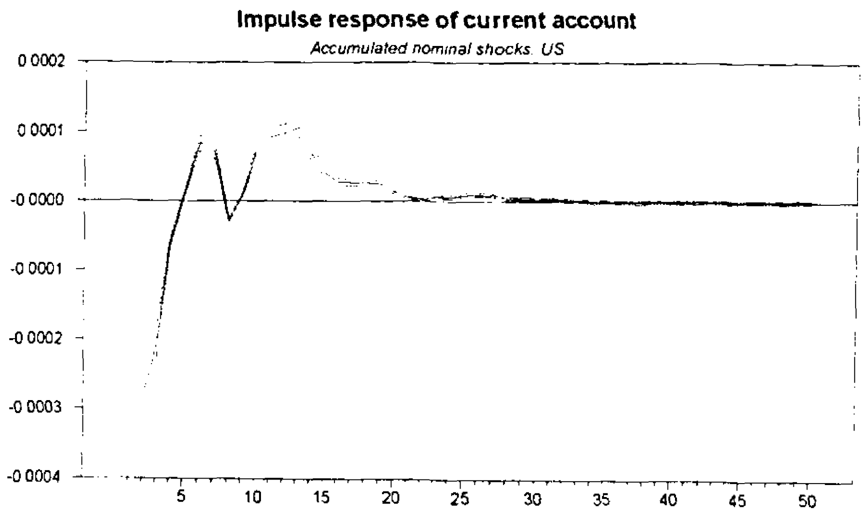


Figure A3.7: United States



3.6.2.2 Extended OECD Sample

Figure A.3.8: Impulse Responses for Austria, Australia, Belgium, Denmark, Finland, an South Korea

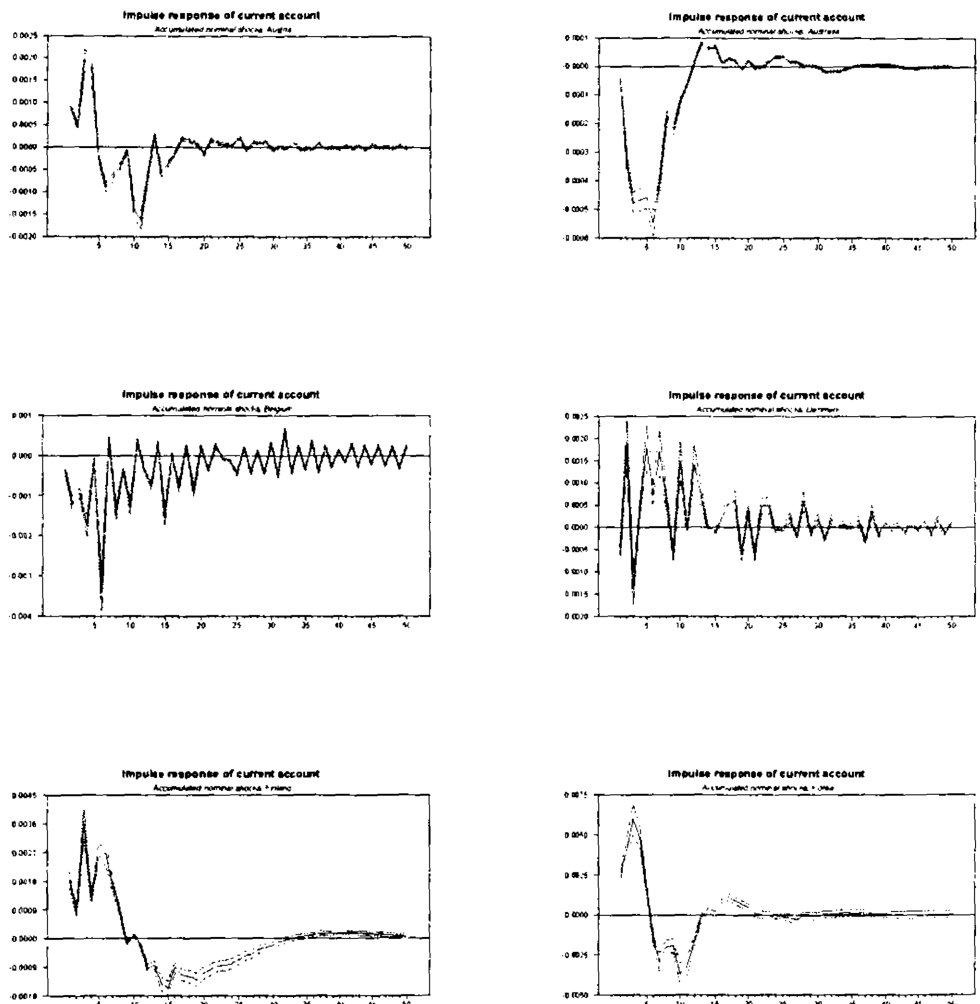


Figure A3.9: Impulse Responses for Mexico, The Netherlands, New Zealand, Portugal, Spain, and Sweden

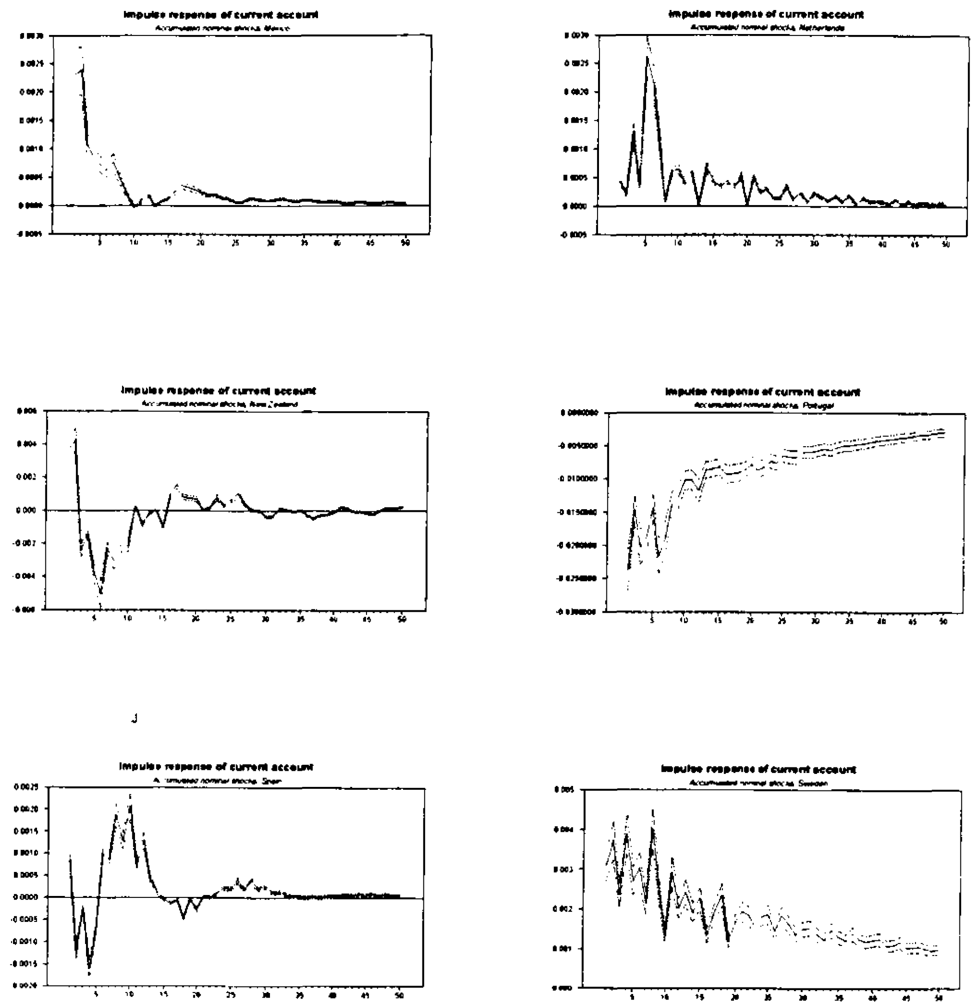
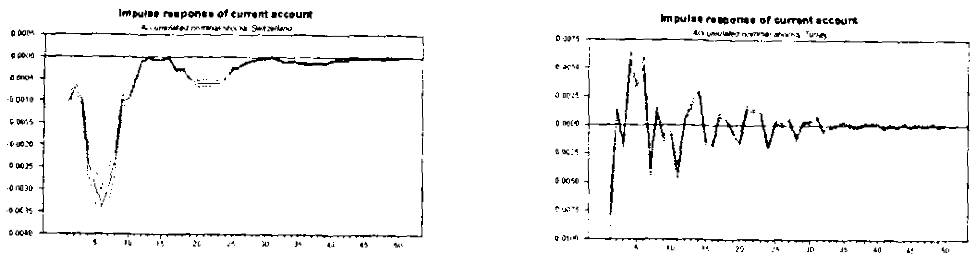


Figure A3.10: Impulse Responses for Switzerland and Turkey



Chapter 4

Is the Output Gap Useful for Economic Policymaking?

> "The quantification of potential output—and the accompanying measure of the 'gap' between actual and potential—is at best an uncertain estimate and not a firm, precise measure."
> (Arthur M. Okun, 1962)

4.1 Introduction

The output gap—which measures the deviation of GDP from its potential—is a frequently-used indicator of the cyclical position and the degree of slack in the economy. Defined as the difference between actual and unobservable potential output, the output gap is itself an unobserved variable. There are numerous ways to calculate potential output, and, consequently, the corresponding output gap. In this chapter,[1] I investigate how to calculate the output gap, whether outcomes are different across methodologies, and whether eventual differences matter. To this end, various measures of the output gap—akin to different methodological approaches to determining

[1] A version of this chapter was circulated previously under Billmeier (2001a).

potential output—are discussed for a small sample of European countries—Finland, France, Greece, Italy, and the United Kingdom.[2]

Information on the cyclical position is important for a number of analytical reasons. First, variations in output, when assessed relative to potential, have distinct implications for inflationary pressures in the economy.[3] Consequently, assessing the output gap is pivotal in the discussion of monetary policy, such as in the Taylor rule or in the inflation targeting (IT) framework (see chapter 2). Second, the cyclical position—as expressed by the size and sign of the output gap—is an important component of calculating the "structural fiscal balance," which aims to gauge the thrust of fiscal policy. Third, the magnitude of the output gap is relevant for assessing economic growth and its components—that is, it helps to evaluate whether variations in actual growth can be attributed to cyclical factors (such as slow growth in trading partner economies) or to a longer-term change in potential growth.

The use (and abuse) of the output gap in policymaking and -prescribing is manifold. In the field of monetary policy, much of the discussion over the last decade or so has focused on the advantages of rules (as opposed to discretion) and all major monetary policy regimes rely—at least implicitly—on some form of the output gap. In a seminal paper, Taylor (1993) proposed a simple instrument rule that tracks U.S. monetary policy surprisingly well during the 1980s and early 1990s. In the simplest form of the rule, interest rates are adjusted according to deviations of inflation from a target level and of output from its trend—that is, the output gap. Since then, uncountable papers have investigated related issues.[4] Although the Taylor rule is a

[2]In other words, this sample of countries represents a relatively heterogenous set of small and large European economies, including the European G-7 economies with the exception of Germany due to data problems.

[3]This is particularly obvious in the definition of potential output contained in Okun (1962): "...the maximum production without inflationary pressure; or, more precisely, a point of balance between more output and greater stability."

[4]Svensson (2003) provides an extensive survey. More specifically, the consequences of output gap uncertainty for the Taylor rule are discussed by Smets (1998) for the United States, and by Ehrmann and Smets (2003) and Eleftheriou (2003) for the Euro area.

simplistic—and fundamentally descriptive—policy framework, the output gap is, in fact, regarded as important in gauging the outlook for price developments in the Euro area and the United States.[5] It has also been recognized that—notwithstanding the terminology—the output gap plays an important role in Inflation Targeting, at least in its "flexible" form, see Svensson (1999). For the Euro area, Gerlach and Svensson (2003) attribute greater importance to the output gap than to money growth as an indicator for future inflation.

Regarding fiscal policy matters, the concept of the output gap has acquired operational—but not legal—status in the Stability and Growth Pact (SGP) of the European Union to calculate annual fiscal balances, and much public attention has been dedicated to the enforcement of the Excessive Deficit Procedure (EDP).[6] In fact, there is a growing discussion on the need to recast the SGP in terms of "structural" instead of "medium-term" fiscal balance; see, for example, Begg *et al.* (2002). The crucial question, therefore, is whether commonly computed output gap measures form a solid basis for calculating structural balances.

From a growth accounting perspective, potential output is a combination of various factors of production, including technological progress/the Solow residual.[7] Since this is not the focus of the chapter, I will not explore this argument further.

The main results of this chapter are that output gap measures can yield very different outcomes, depending on the method used to determine potential output and that care should be exercised when dealing with output gap measures—and devising policy recommendations based on them. Moreover, there appears to be no *a priori* reason to prefer one measure over another (absent methodological issues, such as

[5]See, for example, ECB (2000) and FRB (2004). For a similar statement regarding monetary policy in New Zealand, see Claus (2000).

[6]The report of the Economic Policy Committee acknowledges the output gap as an essential—but so far only intermediate—input for assessing the progress made by member countries towards achieving the goal of medium-term fiscal balance; see European Commission (2001). Moreover, the report specifies that the production function approach be the one used for policy assessments by the Commission.

[7]See Solow (1956).

inconsistency of a particular method with economic priors).

As a corollary, and to gain further insight, an econometric evaluation is carried out by a simulated out-of-sample forecasting exercise in a Phillips curve framework, which relates the inflation rate to the output gap. If the output gap, in fact, measures domestic inflationary pressures, then a simple autoregressive inflation forecast could be improved by taking into account the information stemming from the output gap. For the forecast period 1990–2002, models are estimated recursively with (real-time) data up to the last year of forecast period—that is, the last observation increases by one year with every iteration—and the one-year-ahead predicted inflation is compared with actual inflation.[8] Based on this evaluation, the output gap is not always a useful measure to gauge domestic inflationary pressures, and no specific measure (in this sample) consistently dominates all other measures or a univariate forecast.

From an empirical perspective, the analysis of output gaps using quarterly (or higher frequency) data would constitute an exercise more consistent with the setup of monetary policy making, because policy decisions are frequent: monthly for the European Central Bank and the Bank of England, and roughly every seven weeks for the Fed's FOMC. However, employing quarterly data comes at a cost that is often overlooked. In fact, data revisions, in particular to quarterly GDP data, are common and sometimes substantial. To minimize the impact of data revisions, I focus in what follows on annual observations.[9] Consequently, the results reported below will be

[8]The evaluation period, 1990-2002, is economically rather interesting in Europe: Finland experienced a major crisis after asset prices collapsed, and Italy and the UK exited from the ERM.

[9]Alternatively, real-time datasets could be analyzed, see Orphanides (2001). Gerlach (2004) offers yet another way to circumvent the data revision problem related to output gaps. He suggests that in line with his reading of the editorial remarks routinely issued by the ECB's Governing Council, frequently-revised output gaps could be substituted by an economic sentiment indicator, which (i) is not subject to revision and (ii) appears to be a leading indicator of two output gap measures in the euro area. While intriguing from a pragmatic perspective, this approach does not tackle another issue addressed in this chapter: the inconsistency of indicators. Different sentiment indicators can be, *a priori*, as dissonant as the gap measures employed in this chapter. Moreover, Orphanides and van Norden (2002) show that ex-post revisions of output gap estimates for quarterly U.S. GNP data are of the same order of magnitude as the gap itself. The bulk of the revisions is attributed

more relevant in the context of fiscal—rather than monetary—policymaking.

The remainder of this chapter is structured as follows. Section 4.2 introduces the literature. Section 4.3 reviews briefly the various output gap measures, and section 4.4 compares them descriptively. Section 4.5 provides the empirical evaluation. Section 4.6 concludes, and section 4.7 contains the appendices to this chapter.

4.2 A Look at the Literature

General research on the output gap started probably with Okun (1962) and has been abundant ever since.[10] Roughly speaking, two broad approaches have been followed in the literature to estimate potential output and the output gap. The first is based on the statistical properties of the underlying GDP series. The second, instead, estimates potential output on the basis of an economic model. As argued in Scacciavillani and Swagel (1999), these different techniques can be viewed as akin to different economic concepts of potential output.

Under the first approach, potential output is driven by productivity shocks, and temporary deviations of actual output result from private agents' decisions to reallocate resources in response to these shocks. Given this (neoclassical) reasoning, potential output coincides with the underlying trend of actual output, and the challenge in estimating the output gap is to separate longer-run changes in the trend from short-lived (temporary) movements around potential. The univariate statistical measures described in Section 4.3 to identify potential output can be traced back to the familiar contribution by Hodrick and Prescott (1997), as well as to Corbae and Ouliaris (2002). The latter authors use frequency domain methods to extract information on the business cycle (and, by implication, the underlying trend) properties

to unreliable end-sample estimates, however, not revisions of published data. Finally, the ex-post correspondence of output gap measures and a sentiment indicator in the euro area may have been an empirical coincidence and may not hold in a different sample.

[10]See, for example, Kuttner (1994) and Cotis *et al.* (*forthcoming*) for reviews.

of GDP.

Under the second approach—somewhat closer to the Keynesian tradition—business cycle swings and hence the gap between actual and potential output reflect demand-determined actual output fluctuating around a slowly moving level of aggregate supply. Thus, any measure of the output gap should account for underemployed resources, in particular in the labor market. This can be done by using an underlying model that describes relevant aspects of the economy. The model-based approaches in the next section relate to Blanchard and Quah (1989), and to the large strand of the literature that focuses on the production function approach to potential output. Contributions include early research done at the IMF such as Artus (1977) and, more recently, De Masi (1997). Proietti, *et al.* (2002) evaluate unobserved components models based on the production function approach for the Euro area as a whole. The implementation of the production function approach follows closely the latent variable approach, developed in Kuttner (1994) and further refined by the European Commission.[11]

While this selection of output gap measures is by no means exhaustive, it brings together some of the most well-known approaches and—given the results—suffices for the purposes of this analysis.[12]

In Section 4.5, the various output gap estimates are compared in an inflation-forecasting framework similar to Stock and Watson (1999), who investigate forecasts of U.S. inflation at the 12-month horizon, using information from 168 additional indicators of economic activity including output gap estimates. In this chapter, however, I focus exclusively on output gap measures and the spirit of the analysis is, therefore, more related to comparative output gap studies such as Scacciavillani and Swagel

[11]See, e.g., Denis *et al.* (2002).

[12]For example, we do not analyze the class of factor-based forecasts; see the series of papers by Stock and Watson (1989, 1991, 1999). Another common decomposition, pioneered by Beveridge and Nelson (1981), is omitted due to insufficient data.

(1999), Cerra and Saxena (2000), Claus, *et al.* (2000), Bolt and van Els (2000), and Ross and Ubide (2001), who, respectively, compare measures for Israel, Sweden, New Zealand, euro area countries, and the euro area as a whole. Recent contributions that evaluate inflation forecasts for the United States and Australia based on real-time estimates of the output gap include Orphanides and van Norden (2002) and Robinson *et al.* (2003). Finally, the statistical test of forecast performance used below is based on Diebold and Mariano (1995).[13]

4.3 Output Gap Measures

The four measures considered decompose a time series (here output, y_t) additively into a cyclical component, y_t^c and a trend component, y_t^*:

$$y_t = y_t^c + y_t^*. \tag{4.1}$$

The trend component is assumed to coincide with potential output, that is the amount of output that can be achieved under normal capacity utilization and given the constraints in the labor market, in particular given the natural rate of unemployment. Accordingly, the output gap measures the relative distance of actual output from trend (the "smoothed" series), that is, the cyclical component:

$$gap_t = \frac{y_t - y_t^*}{y_t^*} = \frac{y_t^c}{y_t^*}. \tag{4.2}$$

The gap measures fall into two broad categories: (univariate) statistical filters, and more theory-related measures based on an underlying economic model.[14] Statistical

[13] An inflation-forecasting framework is clearly not the only way to compare output gap measures. Alternatively, one could ask whether they have predictive value for certain types of tax revenues or cyclical expenditure items. Moreover, the gap should probably correlate with firms' pricing/mark-up strategy. While interesting, these hypotheses are not explored in this chapter.

[14] The output gap measures considered differ is the amount of parameters estimated. The issue of parameter instability due to the relatively short estimation period is acknowledged, but not pursued

filters can extract information either in the conventional time domain—exemplified by the version of the Hodrick-Prescott (1997) filter used below—or in the frequency domain. The frequency domain approach is represented by a filtering method recently developed by Corbae and Ouliaris (2002), drawing on earlier results by Corbae *et al.* (2002).

Among the many measures of potential output that rely to a larger extent on economic theory, this chapter evaluates the permanent-transitory decomposition by Blanchard and Quah (1989) and the production function approach. In the case of the Blanchard-Quah routine, the economic reasoning is tied to the conventional distinction of "demand" versus "supply" shocks, whereas the production function methodology is based on a model of the aggregate production structure of the economy.

4.3.1 The Hodrick-Prescott Filter

The Hodrick-Prescott (HP) filter is probably the most well-known and most widely used statistical filter to obtain a smooth estimate of the long-term trend component of a macroeconomic series. This is chiefly due to its simplicity, but also to the fact that, for the United States, business cycle movements can be extracted using this filter that resemble the NBER-backed definitions (see Canova (1999)). The HP filter is a linear, two-sided filter that computes the smoothed series by minimizing the squared distance between trend (y_t^*) and the actual series (y_t), subject to a penalty on the second difference of the smoothed series:

$$\underset{y_t^*}{Min}\left\{\sum_{t=1}^{T}(y_t-y_t^*)^2+\lambda\sum_{t=2}^{T-1}\left[\left(y_{t+1}^*-y_t^*\right)-\left(y_t^*-y_{t-1}^*\right)\right]^2\right\} \qquad (4.3)$$

The penalty parameter, λ, controls the smoothness of the series by setting the ratio of the variance of the cyclical component and the variation in the second difference

further.

of the actual series. A higher value for λ implies a smoother trend (and, hence, more volatile gaps). In the extreme case of $\lambda \rightarrow \infty$, the trend is a straight line. The standard value in the literature is $\lambda = 100$ for annual data, which is also assumed as a base case in what follows.[15]

In a policy-related context, the traditional Hodrick-Prescott measure poses a substantial problem: the filter as described above is fundamentally a two-sided filter, that is, computation of the underlying trend at time t is based on observations *before* and *after* period t. Economic policymakers, instead, will—at the time of decision-making—only dispose of an estimate of the output gap that is based on a purely backward-looking evaluation of potential output. To avoid this inconsistency, this chapter uses a "real-time" output gap series, *HP.rt*, that is constructed based on the conventional HP filter. This new series consists of "last observations," that is, real-time estimates of the underlying trend in the last observation period t given the information set in period t.

This way to proceed is subject to two important caveats. First, an observation for output produced in period t has to be a prediction while the economy is still in period t and will be finally observed only in $t+1$ (for example, data on 1999 GDP is only issued (at best) in the course of 2000). Second, data may be revised in later periods. I abstract from these important details, since the empirical analysis will build on yearly observations, implying that by the end of a given year, the first three quarters of the yearly figure for output have already been observed and provide a sound footing for an end-year estimate. Annual data are also less likely to be affected by substantial data revisions since these revisions usually occur in the periods immediately following the (quarterly) observation, hence, mostly before the end of the calendar year. In addition, revisions due to seasonal factors are limited for annual data.

[15]This follows Hodrick and Prescott (1997) and was already implicitly argued by Burns and Mitchell (1946); see Section 4.5.3.1 for a robustness check regarding the value of λ. Ross and Ubide (2001) discuss alternative approaches to determine the parameter λ endogenously.

Other prominent drawbacks of the HP filter (in the version described above) have been documented in the literature and include the possibility of finding spurious cycles for integrated series, the somewhat arbitrary choice of λ, as well as the neglect of structural breaks and shifts.[16]

4.3.2 The Frequency Domain Approach

Economic fluctuations occur at different frequencies (displaying, for instance, seasonal, or business cycle duration). Starting from the classical assumption contained in Burns and Mitchell (1946) that the duration of business cycles takes between 6 and 32 quarters, the approach to extracting those cycles from a stationary time series is relatively straightforward from the frequency domain perspective. The original series should be filtered in such a way that fluctuations below or above a certain frequency are eliminated.

This can be achieved with the help of an exact band-pass filter (BPF). An exact BPF acts in principle as a double filter: it eliminates frequencies outside (that is below and above) a specified range, here the business cycle frequency. For estimation purposes, however, these filters are usually spelled out in the time domain, since integrated series—such as real GDP—could traditionally not be handled by frequency domain approaches.[17] However, transformation of the exact band-pass filter back into the time domain results in a moving average process of infinite order. For this reason, Baxter and King (1999) and others have provided time domain approximations to

[16] See, e.g., Harvey and Jaeger (1993), King and Rebelo (1993), and Cogley and Nason (1995) for overviews of the shortcomings. Billmeier (2004b) provides an illustration of another problem of the HP filter, the end-sample bias. The discussion of the optimal λ is circumvented here by comparing three values, see section 4.5.3.1. Ravn and Uhlig (2002) argue a value of 6.25 for annual observations, based on the assumption that $\lambda = 1600$ is the optimal value for quarterly data (which is not necessarily true for our sample). Artis *et al.* (2002) argue the superiority of the band-pass version of the HP-filter.

[17] As Corbae and Ouliaris (2002) explain, this is due to a "leakage problem:" the frequency responses generated by the discrete Fourier transform of an I(1) process are dependent across fundamental frequencies.

the exact band pass filter capable of dealing with integrated series. Their method involves a trade-off between the quality of approximation and the ability to smooth the series at the extreme points of the sample, since every additional lag employed in the estimation process improves the filter but translates into one lost observation at either end of the series. This, in turn, substantially diminishes the attractiveness of this class of filters for policy-related analysis. Alternatively, the estimation can take place directly in the frequency domain. According to Baxter and King, pre-filtering of the non-stationary series is required to remove stochastic trends and, hence, avoid the leakage problem with integrated series. They argue that upfront detrending of the series in order to apply discrete Fourier transforms involves a discretionary choice of the detrending method, whereas the symmetric moving average approximation would successfully remove any deterministic or stochastic trends up to second order.

In this context, Corbae and Ouliaris (2002) provide a frequency domain fix for the leakage problem and, hence, a consistent band pass filter for non-stationary data. In addition, the filter does not involve a loss of observations at either end—a property highly relevant for policymaking.[18] In the base case econometric evaluation in Section 4.5.2, a business cycle duration between 2 and 8 years is assumed.[19]

While the major advantage of the frequency domain approach and, indeed, other statistical methods not mentioned here such as arithmetic detrending is their simplicity, they are subject to the criticism of lacking foundation in economic theory. Thus, the next two sections turn to theory-based models of trend GDP and the output gap.

4.3.3 The Blanchard-Quah Decomposition

The appeal of the approach by Blanchard and Quah (1989) to the identification of structural shocks in a VAR stems from its compatibility with a wide array of

[18]See Corbae and Ouliaris (2002) for a technical description of the filter and its small sample properties, and Corbae *et al.* (2002) for the analysis of the asymptotic case.

[19]Robustness checks are contained in section 4.5.3.1.

theoretical models. In a bivariate system, structural supply and demand shocks are identified by assuming that the former have a permanent impact on output, while the latter can only have a temporary effect. In particular, two types of (uncorrelated) structural disturbances are postulated, which possibly affect two time series, (log) real GDP and the unemployment rate. The following assumptions identify these disturbances: no disturbance has long-run effects on the time series employed in the estimation, more precisely on the first differences of the original time series (i.e., growth rates are stationary). Furthermore, disturbances to (the growth rate of) real GDP may have long-run effects on the level of both series, while disturbances to the unemployment rate are restricted to not having long-run effects on the level of output. These assumptions technically identify the shocks. Given the chosen structure, it seems natural to label the shocks as supply and demand shocks.[20]

In the present context, potential output is associated with cumulated supply shocks, whereas the output gap reflects cyclical (temporary) swings in aggregate demand. This approach, hence, benefits from explicit economic foundations. Furthermore, the gap—identified as the demand component of output—is not subject to any end-sample bias. However, the identification scheme employed may not be appropriate under all circumstances, in particular if the variable representing demand (here the unemployment rate) does not provide a good indication of the cyclical behavior of output. Finally, given the orthogonality assumption on the structural shocks, the amount of variables also determines the number of shocks present in the system. Conversely, there are clearly shocks that have a supply as well as a demand component, for instance, public infrastructure investment.

The VAR models estimated include, in addition to a constant, up to four lags of the endogenous variables, as indicated by information criteria. No residual autocorrelation was present in the specifications chosen.

[20]The empirical set-up has been documented numerous times in the literature, see, e.g., the previous chapter and the literature cited therein for a more detailed description of the approach in the context of a three-variable model.

4.3.4 The Production Function Approach

A way to circumvent the problem of assigning shocks to demand or supply origins is to start from a growth-accounting perspective. The production function approach describes a functional relationship between output and factor inputs. While the method as such is not new, recent focus on the input factors, in particular labor, has triggered new interest in the subject. I describe both issues in turn.

4.3.4.1 The Input-Output Relationship

Output is at its potential if the rates of capacity utilization are normal; that is, labor input is consistent with the natural rate of unemployment and technological progress/total factor productivity is at its trend level. A convenient functional form is the Cobb-Douglas type, where output Y_t depends on labor L_t and capital K_t, as well as the level of total factor productivity TFP_t:

$$Y_t = TFP_t K_t^{\alpha} L_t^{\beta}. \tag{4.4}$$

Assuming constant returns to scale implies that $\alpha + \beta = 1$; under perfect competition, α corresponds to the share of capital income, and $\beta = 1 - \alpha$ to the share of labor. Since total factor productivity is not observable, it is usually derived as a residual from the above equation:

$$tfp_t = y_t - \alpha k_t - (1 - \alpha) l_t \tag{4.5}$$

where variables in small caps are in logs. Log trend TFP, tfp^*, is then obtained by appropriately smoothing this residual series, for instance by a HP filter. Potential labor input L^* is taken to be the level of employment consistent with the (time-varying) natural rate of unemployment UR^*:

$$L_t^* = LF_t\,(1 - UR_t^*) \tag{4.6}$$

where LF_t is the labor force. Potential output can be written (in logs) as:

$$y_t^* = \alpha k_t + (1-\alpha) l_t^* + tfp_t^*. \qquad (4.7)$$

The most important advantage of the production function approach lies in its tractability together with the possibility to account explicitly for different sources of growth. For instance, the dynamic growth of the Finnish ICT sector during the second half of the 1990s had been mostly driven by potential growth from a productivity point of view and, hence, had resulted in a rather small output gap (see Section 4.4.2). Moreover, the strong movements of the unemployment rate since the Finnish crisis in the early 1990s convey valuable information on labor market conditions. Important shortcomings of the approach include the dependence on a number of crucial assumptions, e.g., (constant) shares of capital and labor, and the functional form of the production relationship (number of input factors, returns to scale). In addition, data requirements can pose significant problems; for example, the capital stock is difficult to measure consistently, in particular at a frequency other than annual.

4.3.4.2 Factor Inputs and the NAWRU

A crucial feature of this approach is the reliance on filtered factor input series, in particular the trend total factor productivity and the natural rate of unemployment. Given the assumption that capital is always employed at full potential, however, no capacity adjustment is usually made to the capital stock.[21] The natural rate of unemployment can be derived in a number of ways, for example by HP-filtering the observed unemployment rate.[22] However, the approach can also be implemented

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[21] In other words, the full-capacity stock of capital is usually approximated by the actual stock of capital. Artus (1977) provides an early attempt to account for capacity usage. For a more recent approach, using French data on capital operating time, see Everaert and Nadal-De Simone (2003).

[22] Of course, the choice of a filter to detrend the unemployment rate and TFP adds an element of discretion.

more flexibly by using more sophisticated filtering procedures, including those that incorporate themselves structural assumptions based on economic theory. Here, the calculation of the output gap using the production function approach emphasizes the derivation of the NAWRU (non-accelerating wage inflation rate of unemployment) as a latent variable following Kuttner (1994).[23] From a conceptual point of view, however, this approach rests on the premise that a natural rate of unemployment exists, in other words that the Phillips curve is vertical at said natural rate. This holds true for the countries in the sample (see appendix 4.7.1).

Under the latent variable approach, the natural rate of unemployment—defined here as the NAWRU—is computed using a Kalman filtering process on the observable unemployment rate to extract the cyclical component. The procedure employs a bivariate model, where the observables "unemployment rate" and "change in wage inflation" (that is, second differences of wages) play the role of endogenous variables. While the first equation contains a simple decomposition of the observed unemployment rate in trend and cyclical component, the second equation—in principle a Phillips curve—relates the wage inflation to a number of regressors, including lags of wage inflation and the cyclical component of unemployment. Given the error term, wage inflation is assumed to follow an ARMA process. The trend unemployment rate, in turn, serves to determine the (full-employment) stock of labor entering the production function. Estimation takes place in the state-space form, some exogenous regressors (such as a variable reflecting terms of trade) are added for some countries to (marginally) improve the statistical fit.[24]

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[23] This approach was recently adopted by the European Commission, see Denis *et al.* (2002) and Planas and Rossi (2003). In the Commission's work, the new methodology substitutes for more "traditional" approaches—such as the Hodrick-Prescott filter—and, at the same time, unifies the Commission's efforts toward a consistent representation of business cycles in the member countries.

[24] A more detailed description of the model set-up can be found in appendix 4.7.2. In the terminology of the European Commission, the NAWRU model is known as the "GAP model." In a related paper, Billmeier (2004b), it is shown that both the assumed representation of wage inflation and the inclusion of additional regressors can have substantial impact on trend unemployment.

4.4 Comparing Output Gaps

In this section, the output gap measures are compared descriptively. The data on real GDP for the five countries in the sample (Finland, France, Greece, Italy, and the United Kingdom) are annual and stem largely from the European Commission's database, as does the unemployment rate representing the demand side-related variable in the Blanchard-Quah decomposition. The same holds true for the variables used to determine the NAWRU and, ultimately, the output gap according to the production function approach. The price indices used in the inflation-forecasting exercise in Sections 4.5.2 and 4.5.3.3 come from the International Monetary Fund's *International Financial Statistics* (IFS) database.[25]

4.4.1 Descriptive Priors

Descriptively, the mean of the gap measure should be close to zero over longer time horizons. Also, an extended period—say, more then 15 years—of expansion or contraction would run counter to the concept of the business cycle *per se*. Finally, the measure should capture a number of stylized facts, in line with traditional descriptions of economic activity in the respective country. In Finland, for example, the measure should reflect the low-inflation boom in the late 1980s and the subsequent overheating; the near-collapse of economic activity during the crisis period 1990–93; and, again, the period of strong economic growth during the late 1990s, driven, at first, by the economic recovery and, later, by the information and communication technology (ICT) boom. In both Italy and the United Kingdom, a major recession preceded the exit from the EMS in the early 1990s. Most approaches—albeit to varying degrees—reproduce these stylized facts.

[25]The CPI corresponds to IFS line 64...ZF, the GDP deflator to line 99BIRZF.

4.4.2 Descriptive Assessment

Table 4.1, left panel, presents descriptive statistics for the four standard output gap estimates considered above (and reproduced graphically in Appendix 4.7.4). In the right panel, a correlation matrix, statistics below the main diagonal represent simple correlations of the main gap measures in levels to gauge the similarity of gap measures. The upper triangle, instead, offers correlations of first differences, which indicate whether the gap measures convey the same directional message, that is, whether the economy is improving or not. A number of observations can be made. First, most gaps—with the possible exception of the one based on the BQ decomposition—are centered around zero. Maxima and minima seem to be in a reasonable range and speak—together with the standard deviation—of the more or less bumpy road the sample countries have travelled down. Particularly interesting in this context is France, where the consistently lowest variation in the sample reflects relatively smooth economic growth, close to potential. The opposite is represented by Finland. Both the extrema and the surprisingly high standard deviation reflect the boom-bust cycle in the early 1990s, when a period of strong expansion ended and an economic downturn, unrivaled among OECD countries after WWII, set in.[26]

Another striking feature of the table is that the 2002 gap estimates are not consistently positive or negative for any single country. This intrinsic uncertainty regarding the output gap is also reflected in the sometimes surprisingly low correlations between the various measures. Across countries, the BQ decomposition seems to yield an output gap measure that is somewhat "distinct" from the other three. This is true for correlations in levels and first differences.

Figures A4.1–A4.3 in the appendix present the output gap measures for the five countries in the sample. The above-mentioned boom-and-bust cycle in Finland in the

[26]See Berger and Billmeier (2003) for a more detailed evaluation of the Finnish experience.

[27]Output gap in 2002.

Table 4.1: Descriptive Output Gap Statistics, 1960–2002

	Sample statistics					Correlations			
	Mean	Min	Max	SD	Gap[27]	HP100rt	PF	BQ	FD2-8
					Finland				
HP100rt	0.08	-7.44	5.77	3.20	-0.25	1	0.94	0.44	0.83
PF	-0.08	-6.49	6.43	3.00	-1.55	0.82	1	0.48	0.87
BQ	-0.38	-9.52	4.00	3.27	1.06	0.09	0.27	1	0.48
FD2-8	0.00	-5.32	5.06	2.61	-1.42	0.57	0.70	0.35	1
					France				
HP100rt	0.07	-2.74	2.81	1.41	0.46	1	0.90	0.18	0.88
PF	-0.13	-2.72	2.81	1.43	0.07	0.81	1	0.01	0.79
BQ	0.22	-1.11	2.70	0.99	0.66	0.36	0.26	1	0.17
FD2-8	0.00	-2.44	2.07	1.01	-0.90	0.66	0.50	0.08	1
					Greece				
HP100rt	0.00	-5.04	5.40	2.36	2.78	1	0.96	-0.05	0.95
PF	0.69	-3.19	4.65	4.83	0.42	0.77	1	0.09	0.91
BQ	-0.43	-8.13	8.59	3.23	-0.12	-0.32	0.12	1	-0.09
FD2-8	-0.04	-3.89	5.19	1.84	0.61	0.68	0.62	-0.12	1
					Italy				
HP100rt	-0.07	-7.12	3.33	1.95	-0.28	1	0.78	-0.17	0.83
PF	-.34	-4.76	3.02	1.85	-0.81	0.66	1	0.32	0.72
BQ	0.96	-5.51	5.54	3.98	0.22	0.11	0.20	1	-0.30
FD2-8	-0.00	-4.14	3.57	1.53	-0.53	0.55	0.50	-0.08	1
					United Kingdom				
HP100rt	0.06	-3.79	5.21	1.86	0.00	1	0.98	0.55	0.92
PF	-0.09	-4.18	5.32	1.99	-0.78	0.92	1	0.47	0.89
BQ	-0.38	-9.32	4.00	3.27	1.06	0.64	0.57	1	0.54
FD2-8	0.06	-3.09	4.57	1.47	0.44	0.70	0.64	0.43	1

Notes: HP100rt, real-time Hodrick Prescott filter; PF, production function approach; BQ, Blanchard-Quah decomposition; FD2-8, frequency domain filter. In the right panel, the lower triangle gives level, the upper triangle first-difference correlations.

early 1990s is clearly visible, as are the causes and consequences of exiting the ERM for Italy and the United Kingdom. For most countries, the dispersion of gap measures seems to increase toward the end of the sample period, with Greece being a particularly good example. The generally low correlation of the Blanchard-Quah-based measure is clearly visible, especially, again, in Greece. The figures and corresponding cross-country correlations (not presented here) also document the role of Greece and, to a lesser extent, the United Kingdom as outliers from the "European" business cycle.[28]

An additional descriptive statistic—measuring the consistency of gap signals—yields broadly similar results (See Table B4.1 in the appendix). This measure is constructed as the share of total observations in which two gap measures give the same cyclical (i.e., boom or bust) signal.[29] Similar to the correlation statistics, the BQ measure again displays the lowest consistency with other measures for most countries.

Summing up the descriptive statistics, it *does matter* how the output gap measure is constructed. In fact, there appears to be a surprising heterogeneity across gaps for a given country. In addition, there is no *a priori* reason to rely on one specific measure as opposed to another. From a policymaking perspective, this is bad news: Diverging quantitative information on the output gap could be taken into account in more gradual policy responses, but if it is not clear whether the gap is positive or negative, few policy recommendations can be derived from such a measure.

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[28] For both countries (and across all gap measures), the degree of business cycle integration did not increase substantially after 1990 when compared to pre-1990 (with the exception of the U.K.-Finnish cycles). While an interesting subject in itself, the issue of business cycle synchronization is beyond the scope of this paper.

[29] This measure differs from the first difference of the gap considered above in that two measures could signal an improvement ($\Delta gap > 0$) but not necessarily the same cyclical position (negative vs. positive gap).

4.5 A Corollary: Using Output Gaps to Forecast Inflation

As a corollary, I use a simple forecasting exercise to gauge the quality of the output gap measures as indicators for inflation. While not intended as a full-fledged inflation-forecasting model, the set-up is common in the literature to evaluate inflation indicators.[30] The approach clearly offers shortcomings compared to more elaborate inflation-forecasting frameworks, in particular those relying on a broader set of variables.[31] In particular, the usefulness of the output gap measures themselves may depend on other information included in the forecasting exercise.[32] As such, this approach cannot be expected to provide the best inflation forecast possible and should, hence, be only taken as one possible way to evaluate the quality of output gaps as indicators for inflation and as an illustration of the results in Section 4.4.

Moreover, due to the policy perspective taken in this chapter, I abstract from one important issue: short-run inflationary pressures—as measured by the output gap—may only be one out of many competing factors that contribute to inflation. Longer-run, supply-side shocks are in principle neglected in the present framework or, to be more precise, affect potential output, that is, the denominator of the gap measure. While of paramount importance, those long-run issues do not take center stage in (monetary) policy circles, which are much more concerned with the short- to medium-run horizon.

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[30]See, e.g., Orphanides (2001), Orphanides and van Norden (2002), Robinson *et al.* (2003), or Stock and Watson (1999).

[31]See Stock and Watson (1999), who find that an aggregate of the 168 time series they have at their disposition is the best indicator for U.S. inflation.

[32]See Chadha and Nolan (2004).

4.5.1 Remarks on the Methodology

At this stage, several remarks about the empirical strategy appear appropriate. First, predictability tests can be based on the in-sample fit of a model or on the out-of-sample fit obtained from a sequence of recursive or rolling regressions. In the context of inflation-forecasting, the latter set-up mimics the data constraints faced by a policymaker in real time, and appears to be, hence, a more appropriate evaluation technique.

Second, the tools used by practitioners for ranking models are the same whether the forecasting models are nested or not. In applied work, forecasting models are often chosen (or dismissed) on the basis of their root (predictive) mean square error (RMSE) compared to a base forecast or a derivative thereof; a well-known example is Meese and Rogoff (1983). However, using relative RMSEs for model selection—as it is also done below—comes at the drawback that there is no straightforward measure of significance.

Third, the output gap is not directly observable and all the standard econometric caveats about two-stage estimation apply. In principle, inference about second-stage inflation forecasts needs to deal with underlying parameter uncertainty of the output gap measure estimates in the first stage; see West (1996). Instead, a test statistic proposed by Diebold and Mariano (1995) has become one of the more common ways to compare alternative forecast paths to the true series—and is also applied further below. This test provides information on whether forecast i is significantly better than forecast j but does not take parameter uncertainty into account. In fact, only two of the gap measures examined, the production function approach and the BQ filter, are affected by first-stage parameter uncertainty. The other two filters (HP and frequency domain) are based on assumptions for the crucial parameters and suffer "only" from the conventional sampling uncertainty. Furthermore, West (1996) shows that under the assumption that OLS provides consistent estimates of the parameters (such as in the VAR used to estimate the BQ decomposition) one can safely ignore parameter uncertainty when testing for differentiable functions of parametric forecasts and forecast errors such as the mean square error.[33] Consequently, the Diebold-Mariano test statistic is applied to shed some further light on the usefulness of the output gap measures chosen in predicting inflation.

Fourth, the forecasting exercise conducted below intends to evaluate the output gaps, but is by no means a full-fledged inflation-forecasting model. It tries to capture the aspect of price developments that is triggered domestically by the cyclical position of the economy; for example, the price adjustment in factor markets according to the factor's marginal rate of utilization. Many other inflation theories are available that capture other variables crucial for price developments, for example the exchange rate.[34] More comprehensive models will undoubtedly lead to firmer results, in particular to a better performance of the forecast relative to the benchmark, the univariate model. For example, factor models, which take into account a much richer variable set, have proven to yield promising outcomes; see Stock and Watson (1999). In the present context however, less emphasis is put on the construction of a well-performing model from a forecasting perspective than on the simple assessment of various output gap measures in an inflation-forecasting context.[35]

4.5.2 Econometric Evaluation

The output gap is often considered a useful instrument to gauge (domestic) inflationary pressures. Consequently, the information stemming from an output gap measure could increase the precision of inflation forecasts. These forecasts, stemming from

[33]The same holds true for a limited number of other cases, including the one of a large estimation sample size relative to the prediction sample size; see, e.g., Diebold (2001). McCracken (2000) points out that under the same conditions, parameter uncertainty is not necessarily irrelevant for the moments of nondifferentiable functions of parametric forecasts and forecast errors such as the mean absolute error.

[34]A substantial amount of research effort has been dedicated to the exchange rate pass-through into domestic prices; see, e.g., Goldberg and Knetter (1997) for a review.

[35]An obvious drawback of factor models in a policy context is the enormous amount of data needed and the sometimes tedious variable handling, see Cotis *et al.* (*forthcoming*).

the gap measures constructed above, are compared using a simulated out-of-sample methodology. The forecasting model is a variant of the Phillips curve:

$$\pi_{t+1} - \pi_t = \alpha + \beta(L)gap_t + \gamma(L)\Delta\pi_t + \epsilon_t \tag{4.8}$$

where π_{t+1} denotes the one-year ahead inflation in the price level at period t, π_t is the actual inflation in period t, gap_t denotes the output gap measure (in levels), L is the lag operator, Δ is the difference operator, and ϵ_t an i.i.d. error.[36,37]

This specification of the forecast equation mirrors a classic Phillips curve, with the output gap measure substituting for the unemployment rate.[38,39] The sample data starts in 1960, the evaluation period spans from 1990 to 2002, and observations are annual. The simulated out-of-sample procedure consists of the following steps. First, a model is estimated for the period 1960 through 1989 with data available up to 1989. Lag length selection of each estimated model up to a maximum number of

[36]This version of the model is chosen following Stock and Watson (1999). Proietti *et al.* (2002) find that the first difference of the output gap (but not the level) is a significant predictor of inflation in the Euro area. Selected experiments have been performed with differenced output gap measures. These experiments yielded results close to the ones presented and have, hence, been omitted; see also the rather similar correlations between gaps in levels and differences in Table 4.1.

[37]The literature on the New Keynesian Phillips Curve (NKPC) has reignited the the dabate over the correct formulation of the Phillips curve; see Galí and Gertler (1999) and Chadha and Nolan (2004). Sbordone (2002) argues that the NKPC is an appropriate framework to predict inflation if the output gaps measures capture well firms' marginal cost.

[38]As Stock and Watson (1999) point out, this specification assumes that (i) the inflation rate is integrated of order one (I(1)); (ii) x_t is I(0); and (iii) both are, hence, not cointegrated. Moreover, the constant intercept implies that the "natural rate" of the output gap is constant. In this literature, inflation is commonly modeled as an I(1) process. Results not reported here have confirmed this assumption for wage and CPI in the sample countries; for Finland see also the discussion in the appendix to this chapter. While the output gap may seem to behave like an integrated process over limited periods of time, it is clearly mean-reverting from a theoretical perspective.

[39]Given that a main ingredient of the output gap based on the production function approach—the natural rate of unemployment—is derived from a similar framework, the evaluation could expected to be biased in favor of this approach. This, however, does not hold true for at least two reasons: (a) the framework described in appendix 4.7.2 is based on wage inflation whereas the evaluation measures performance in forecasting CPI inflation; and (b) the natural rate of unemployment is only one building block of potential output according to the production function approach—with total factor productivity being quantitatively much more important most of the time.

lags is based on minimization of the Akaike information criterion. Due to the low frequency of the data and the limited number of observations, a maximum of two lags is chosen.[40] Second, a one-year-ahead forecast for inflation in 1990 is made. Third, this value is compared to the actual inflation, yielding the forecast error. Next, the same procedure is repeated including data until 1990. This exercise is computed recursively, that is, for every year until 2001, the model is re-estimated according to the new information criteria, and the forecast error is computed. This procedure yields a unique series of forecast errors for each output gap measure considered.[41]

Table 4.2 presents two statistics: in addition to the cumulative root mean square error (RMSE), the so-called U statistic proposed by Theil (1971) is given. The latter consists of the RMSE of a specific inflation forecast standardized by the RMSE of the naïve forecast of "no change" (NC) in inflation. A value smaller than unity stands for a smaller RMSE than under the naïve hypothesis. The obvious advantage of this statistic lies in its ease of comparability across countries.[42] As additional—and often more challenging—benchmarks, the models were also estimated without any output gap measures, that is, as a univariate inflation forecast ("AR") and under the classic Phillips curve specification, which includes the unemployment rate as a RHS variable ("UR").[43] In the table, specifications that "beat" the univariate model, a very common benchmark in the literature, are in bold.

The results can be summarized as follows:

[40] Section 4.5.3.2 examines the robustness of this assumption.

[41] Parameter estimates of equation (8), while not reported to conserve space, are broadly in line with expectations. In particular, $\widehat{\beta}$ is estimated to be positive and frequently significant.

[42] In other words, cross-country comparison is not biased by the quality of the benchmark. Using a different comparator, say, the autoregressive forecast, statements such as "gap x performs Y percent better than the benchmark in country z" would depend on the quality of the benchmark forecast. The "quality" of the no-change forecast depends solely on the volatility of the series itself.

[43] The order in the autoregressive forecast and number of lags in the classic Phillips curve specification is given by the maximum number of lags used in the model estimation; that is, two in the base case analyzed in this section. See also section 4.5.3.2.

Table 4.2: Evaluation of Forecast Performance I

	HP100rt	PF	BQ	FD2-8	NC	AR	UR
			Finland				
RMSE	1.90	1.71	1.18	1.52		1.14	2.22
Theil's U	1.24	1.11	0.77	0.99	1	0.74	1.45
			France				
RMSE	0.64	1.24	0.55	0.54		0.53	0.78
Theil's U	0.93	1.81	0.80	0.78	1	0.77	1.13
			Greece				
RMSE	2.93	**1.99**	3.24	2.50		2.23	**2.07**
Theil's U	1.20	**0.81**	1.33	1.02	1	0.91	**0.85**
			Italy				
RMSE	1.11	1.00	**0.87**	0.96		0.93	0.94
Theil's U	0.98	0.88	**0.77**	0.85	1	0.82	0.84
			United Kingdom				
RMSE	**1.53**	2.06	1.99	1.84		1.73	1.87
Theil's U	**0.67**	0.91	0.87	0.81	1	0.76	0.82

Notes: see Table 4.1; NC, assumption of no change; AR, autoregressive estimate; UR, unemployment rate (Phillips curve specification).

- No output gap measure yields better results than a univariate inflation forecast in Finland and France. In Finland, the disappointing performance of the output gaps in improving the inflation forecast is most likely due to the high volatility of output itself (see Table 4.1), hampering the determination of a statistically satisfying measure of potential output.

- In both Finland and France, the classic Phillips curve featuring the unemployment rate fares much worse than the naïve forecast, whereas this does not hold true to the same extent in the other three countries. This indicates that in these countries, the unemployment rate is not a particularly good indicator for inflationary pressures stemming from the labor market. In fact, large-scale active labor market policies reduce the amount of *de facto* unemployed in Finland by up to 40 percent, see Feldman *et al.* (2003).

- In Greece, adding either a production-function-based output gap measure ("PF") or the unemployment rate (i.e., the classic Phillips curve specification) improve the inflation forecast performance.

- In Italy, the gap measure based on the Blanchard-Quah decomposition ("BQ") and particularly the frequency domain approach assuming a business cycle length between two and eight years ("FD2-8") seem to provide a better grip on the data than the univariate model.

- The United Kingdom is the only country in the sample where the very popular HP filter in its real-time version for $\lambda = 100$ ("HP100rt," together with the Blanchard-Quah decomposition) delivers good results.

Taken together, these results imply that (i) the univariate inflation forecast is not necessarily improved by adding an output gap measure and that (ii) if a forecast-improving output gap exists, it usually varies by country.[44]

To explore the significance of these results, the inflation predictions are assessed using a simple version of the test proposed by Diebold and Mariano (1995). This test assesses whether—conditional on the true series—the inflation path predicted by adding a measure of the output gap is significantly better than the benchmark autoregressive forecast. Consider the original inflation series, $\{\pi\}_t^T$, two forecasts, $\{\pi_t^i\}_t^T$ *and* $\{\pi_t^j\}_t^T$, and the corresponding errors, $\{e_t^i\}_t^T$ and $\{e_t^j\}_t^T$. The loss function associated with a specific forecast i, $l(\pi_t, \pi_t^i)$, will be in many—but not all—cases a direct function of the forecast error, $l(e_t^i)$. The null hypothesis of equal forecast accuracy for the two forecasts can then be expressed as:

$$E[d_t] = 0, \tag{4.9}$$

where $d \equiv [l(e_t^i) - l(e_t]$ is the loss differential. In other words, under the null, the population mean of the loss-differential series is zero. Under the alternative, forecast i is better than forecast j. Empirically, the Diebold-Mariano statistic is simply the

[44] Chadha and Nolan (2004) provide similar evidence on the weak performance of a traditional Phillips curve framework to forecast inflation and suggest to augment the framework by explicitly modelling factor markets.

"t-statistic" of a regression of d on a constant with heteroskedasticity autocorrelation consistent (HAC) standard errors; see Diebold and Mariano (1995) for details.

Table 4.3 presents the Diebold-Mariano (DM) test statistics (and the corresponding p-values) for the forecasts to be equally accurate. The loss function is specified as

$$d = \frac{1}{T}\sum_{n=t}^{T}\left[l\left(e^{i}_{t,T}\right) - l\left(e^{j}_{t,T}\right)\right]. \tag{4.10}$$

In other words, d is the sample mean loss differential. As the benchmark forecast, the autoregressive inflation forecast is considered (as opposed to Table 4.2, in which the base case is "no change"). Failure to reject H_0 implies that the inclusion of the output gap measure does not provide a better inflation forecast than the simple univariate model.

The results obtained by applying the Diebold-Mariano procedure mirror those anticipated in Table 4.2—but go beyond them in one important aspect. While the RMSE ratio indicated a few output gap-based forecasts that could beat the autoregressive forecast in Greece, Italy, and the United Kingdom, the DM statistic reveals that none of these forecasts is *significantly* better than the AR process. In fact, the null hypothesis of similar forecast prediction accuracy cannot be dismissed for any gap-based inflation forecast at conventional significance levels (of 5 or 10 percent).

There are at least two simple explanations for the lack of significance of the estimates. First, domestic inflationary pressures (as captured by the output gap) may not be the main driver of CPI inflation. This observation is particularly relevant for relatively open economies that do not peg their exchange rate to their major trading partners and that are, therefore, more directly affected by fluctuations in the exchange rate.[45] In the sample, this holds true for the United Kingdom (belonging

[45] This is true under the assumptions that the relevant price index, e.g., the CPI, contains imported

Table 4.3: Evaluation of Forecast Performance II

	HP100rt	PF	BQ	FD2-8	UR
		Finland			
DM statistic	2.10	1.51	0.49	2.79	2.36
(p-value)	(0.98)	(0.93)	(0.69)	(0.99)	(0.99)
		France			
DM statistic	0.66	2.87	0.89	0.18	1.91
(p-value)	(0.75)	(0.99)	(0.81)	(0.57)	(0.97)
		Greece			
DM statistic	1.92	-0.81	1.32	0.77	-0.27
(p-value)	(0.97)	(0.21)	(0.91)	(0.78)	(0.39)
		Italy			
DM statistic	1.89	2.13	-0.90	0.26	0.15
(p-value)	(0.97)	(0.98)	(0.18)	(0.60)	(0.56)
		United Kingdom			
DM statistic	-0.56	1.13	0.94	0.39	0.55
(p-value)	(0.29)	(0.87)	(0.83)	(0.65)	(0.71)

more to the U.S. cycle), and, to some extent, also for Finland and Greece. Whereas Finland's trade patterns shifted toward Europe only after the implosion of the Soviet Union, Greece is trading less and less with EU countries—while slowly becoming a hub for the countries in the Balkan. For Italy and France, however, this explanation does not prove satisfactory because both countries' major trading partners were, at least for most of the sample period, linked to Italy and France through a stable peg. Second, the analysis suffers from a lack of observations. Basing the forecasts on annual observations is beneficial in that some of the data revision issues are avoided, but it also comes at the cost of a significantly reduced sample.[46] While the alternative approach—basing the forecasts on quarterly observations—would probably be more relevant from a (monetary) policymaker's perspective, the lack of significance using annual observations clearly shows that the output gaps used in the prediction exercise do not fully capture the business cycle—at least not from the perspective of

goods and that the exchange rate pass-through is positive.

[46]This effect would be even more noticable if the forecast took estimator uncertainty into account when using a different criterion, e.g., the mean absolute error; see section 4.5.1.

inflationary pressures arising in an overheating economy.

From a fiscal policy perspective, these results also cast doubt more broadly on using output gaps to calculate a fiscal balance corrected for the cyclical position of the economy. First, none of the gap measures (which are arguably an indicator of the cyclical position of the economy) seems to capture domestic inflationary pressures well. Second, it is not *a priori* clear which measure of the output gap should be used since there are many and they would result in widely different estimates of an "adjusted" deficit (or surplus).

4.5.3 Robustness Checks

In this section, the results presented above are checked for robustness in three different dimensions. First, for those gap measures that are subject to a crucial single assumption in the first stage, this assumption is modified. This concerns the penalty parameter λ used in the HP filter and the assumption on the business cycle length in the frequency domain filter. Second, the results are tested with respect to a modelling choice in the second step, namely the maximum number of lags in the prediction model. Finally, to minimize the effects of exchange rate fluctuations and changes in indirect taxation on the inflation measure, the consumer price index is substituted with the GDP deflator.

4.5.3.1 Output Gap Parameters

Generally speaking, HP filters based on differing assumptions on λ find a qualitatively similar pattern of the output gap, that is, their turning points coincide chronologically. The assumption on the smoothness of the trend has, however, strong implications for the magnitude of the gap, in particular at the end of the observation period. Since the "real-time" constructs in the previous section consist, after 1990, of a series of "last observations," the choice of λ becomes even more critical. Here, the exercise is repeated for $\lambda = 20$ ("HP20rt") and $\lambda = 200$ ("HP200rt"), generating a more volatile (HP200rt) and a less volatile (HP20rt) gap measure. Moreover, I flag for comparison the results for the traditional (two-sided) HP filter ("HP100," "HP20," "HP200") to gauge the importance of the "real-time" construct. A similar argument applies to the frequency domain filter. To corroborate the above results, the filter has been re-estimated assuming business cycle durations between two and six years ("FD2-6") and between two and ten years ("FD2-10"). To increase comparability with the results described above, the maximum lag length has been held constant (at two).

Results for the alternative parametrization at the first stage (reported in Table B4.2 in the Appendix) confirm to a large extent the outcome obtained above. Again, the estimated output gaps do not improve on the univariate prediction for both Finland and France. For Greece, only one parametrization of the two-sided HP filter (HP200) produces a slightly but insignificantly smaller RMSE than the autoregressive forecast. For Italy and the United Kingdom, the results from varying the parameters are somewhat more encouraging. For Italy, the output gap measure improves the inflation forecast if a shorter cycle (between two and six years) is assumed in the frequency domain approach. More importantly, this version of the frequency domain is the only forecast that is statistically significant at the 5 percent level in the Diebold-Mariano sense, that is, with a test statistic of $d = -1.97$, it significantly improves upon the autoregressive forecast (p-value 0.02). For the United Kingdom, instead, the real-time HP filter seems to be the best way to capture inflationary pressures—almost independently of the assumption regarding the trend smoothness (λ). In fact, both alternative real-time specifications produce smaller RMSEs that the autoregressive forecast. In the Diebold-Mariano test, however, none of the U.K. real-time measure is statistically significant at conventional levels.

An interesting corollary derives from the comparison of the real-time HP filter with its common, two-sided version. The additional information on the relative position in the cycle stemming from future observations could be expected to improve the quality of the gap measure, and hence the forecast performance. This conjecture is rejected for the countries in the sample. For all countries but one (Greece), the real-time versions of the HP-filtered gaps provide better forecasts. This observation is both intriguing and reassuring. Consider the following situation: a country has a consistently positive output gap, coinciding with elevated inflation for an extended period with the exception of one year in the middle of the sample, when, due to an exogenous shock, the output gap is negative, resulting in somewhat lower-than-average inflation in that period and the following one. The two-sided HP filter, in principle, smooths over the outlier, and predicts, consequently, relatively high inflation. With the forward-looking information missing, instead, the real-time filter picks up the drop in inflation to a larger extent, resulting in a reduced smoothing effect and, hence, a more accurate prediction of inflation. From a policy making perspective, this is reassuring since the two-sided HP filter is, of course, not available (or based on forecasts). In Greece, though, the two-sided filter yields better inflation forecasts than the real-time variant. Both are dominated, however, by another gap measure, namely the one stemming from the production function (in addition to the Phillips curve prediction using the unemployment rate).

4.5.3.2 Forecasting Model Parameters

The second-stage robustness check is related to the setup of the forecasting model. Determination of the optimal lag length for the recursive inflation prediction models requires the specification of a maximum number of lags of the RHS variables in equation (8), that is, the output gap and inflation. With the total number of observations being rather small due to the annual frequency of the data, this limit was set at two in the previous section. Changing this assumption can have a strong impact on the degrees of freedom, and implicitly, on the precision of the estimates.

Most of the results obtained above for a maximum of two lags are robust to a higher or lower limit, see Tables B4.3-B4.5 in the appendix. In particular, for all models evaluated (between one and four lags), no output gap measure yields better results than a univariate inflation forecast in Finland and France.[47] In Greece, both the production-function-based output gap measure and the classic Phillips curve specification improve the inflation forecast performance at all levels—but not significantly so.[48] In addition, if a maximum of one lag is allowed, the BQ filter produces an inflation forecast that is significantly better than the autoregressive estimate at the 10-percent level. In all countries but Greece, instead, the unemployment rate does not provide any useful information when compared to the univariate specification, and hence, is not a good indicator for domestic inflationary pressures. Again, the performance of the unemployment rate is particularly disappointing in Finland and France. The results are the least robust for Italy. While for short lag limits (one and two), the frequency domain and the Blanchard-Quah measure yield the best results, the production function approach (and the real-time HP filter) prevail with a maximum of three and four lags. At three lags, the gap measure based on the production function provides a significant improvement over the autoregressive inflation forecast. For U.K. data, the HP filter performs well for three and four lags—but remains insignificant.

4.5.3.3 Inflation Measure

If the consumer basket measured by the consumer price index contains imported goods, the price of these goods is likely to change to some extent if the exchange rate fluctuates—an effect not related to domestic price pressures and, hence, not captured by measures of the output gap.[49] A similar argument can be made for changes in indirect taxation. They are not caused by domestic economic developments, but have

[47]For a maximum of one lag, the frequency domain approach is just as good as the univariate forecast.

[48]The production function approach is almost significant for a maximum of one and three lags, indicating some scope for a better performance of a refined gap estimate based on the production function.

[49]This obviously depends on the degree of exchange rate pass-through. For the sake of the argument, it is sufficient if the pass-though is non-zero.

an impact on the CPI level. To isolate these effects, the forecasting exercise was repeated with the GDP deflator instead of the CPI.

Compared to CPI inflation, results for the GDP deflator are equally disappointing regarding the basic set of output gap measures (see Table B4.6 in the appendix). In terms of the RMSE, some output gap measures produce an inflation forecast that is marginally better that the univariate forecast (for France, Greece, and the United Kingdom), but no measure succeeds in establishing a significantly better forecast according to the Diebold-Mariano statistic.

4.6 Concluding Remarks

In this chapter, I compared a number of commonly used output gap measures in an inflation-forecasting exercise for a small set of European countries. The measures evaluated included variants of the HP filter, the Blanchard-Quah decomposition, the production function approach, and a frequency domain filter. Reflecting domestic inflationary pressures, the unobservable output gap could, at least in theory, provide some information for one-year-ahead actual inflation and, hence, improve a univariate forecast.

So is the output gap a good indicator for inflationary pressures? Judging from the sample countries Finland, France, Greece, Italy, and United Kingdom, the answer is clearly "it depends." In some countries (Finland, France), the widely used output gap measures evaluated in this chapter do not provide any improvement over a univariate inflation forecast. In other countries, however, inflation forecasts are better if some measure of the gap is included (Greece, Italy, United Kingdom). The best measure to be included varies by country, but is, where applicable, robust to alternative ways of computing the measure (Italy: frequency domain method; United Kingdom: HP filter), and modeling assumptions in the inflation forecast exercise (with the possible exception of Italy). Unfortunately, few of these forecasts perform significantly better than the autoregressive forecast in a statistical sense as documented by the Diebold-

Mariano test. From a policymaking perspective, the conclusions are that (i) various measures of the output gap should be taken into account when assessing the cyclical position of the economy; (ii) it is hard to significantly improve on an autoregressive inflation forecast using the simple output gap measures presented above; (iii) a broader set of indicators may be needed to capture (domestic) inflationary pressures well; and (iv) assessing the fiscal stance based on a structural balance is tricky if the gap measure is to reflect the business cycle well.

The failure to identify output gap measures that improve consistently and significantly upon the univariate forecast in all countries—but, in particular, in Finland and France—could also be due to the limited number of observations, given the choice of annual frequency. In principle, this choice is well-motivated by seasonality issues, data revision considerations, and, to some extent, by the availability of data (measures of the capital stock and the natural rate of unemployment in the production function approach). Nevertheless, a quarterly assessment may (where possible) yield additional insights.

Thus, directions for future research in a similar empirical model include: analyzing real-time data series at a quarterly frequency to mimic better the policymaker's decision-making process; extending the analysis to a larger sample of countries; and including "optimized" versions of the HP and the frequency domain filters—as opposed to the "conventional" robustness analysis with regard to the crucial parameters undertaken above. Further insight could be gained by exploring competing output gap measures in a different set-up. From a fiscal perspective, for example, the output gap could help predict certain cyclical expenditure items. Moreover, the New Keynesian literature has spent considerable energy on finding a suitable measure for real economic activity; exploring the robustness of some of these models with regard to competing empirical specifications of the output gap could lead to interesting results.[50]

[50]See Galí and Gertler (1999).

4.7 Appendices

4.7.1 Is There a Long-Run Phillips Curve?

A modeling framework based on a (time-varying) NAWRU—understood as the natural rate of unemployment underlying the economy—implicitly assumes that the Phillips curve is vertical at said natural rate, i.e., the unemployment rate is independent of (wage) inflation. In other words, empirical inference along these lines rules out the existence of a long-run non-vertical Phillips curve, and, hence, an underlying long-run relationship between inflation and the unemployment rate. This prior has been questioned recently by a number of authors (see, for instance, Beyer and Farmer (2002), and Schreiber and Wolters (2003)), who found empirical evidence against the vertical Phillips curve assumption in U.S. and German data, respectively. As the latter argue, the existence of the NAWRU can be rejected if both the unemployment rate and the rate of (wage) inflation are non-stationary and cointegrated, indicating a long-run relationship, similar to a Phillips curve. In results not reported here, I have not found strong indications against the verticality assumption. I demonstrate the approach for the case of Finland, the other countries in the sample yield similar outcomes.

In the case of Finland, simple tests indicate that, while both wage inflation (*dwage*) and the unemployment rate (*ur*) are non-stationary, there is no sign of cointegration, a result conducive to the NAWRU approach.[51] In fitting a bivariate VAR to the basic data, the lag length was chosen according to the Schwarz and the Hannan-Quinn information criteria, which both propose three lags in levels (Table A1.1).[52]

Due to the lack of strong priors in favor of a trend restricted to the cointegrating space, a VAR system with an unrestricted constant was estimated. Likelihood ratio

[51]Note that these conclusions also hold for CPI inflation (results not reported here).

[52]With three lags, no significant residual autocorrelation emerged, whereas more parsimonious models reveal problems of autocorrelation at the first lag.

Table A4.1: Lag Length Selection

lags	Schwarz	Hannan-Quinn
4	-6.0	-6.5
3	**-6.4**	**-6.8**
2	-6.2	-6.5
1	-5.9	-6.0

tests of the time series properties (distributed as $\chi^2(dgf)$, where dgf stands for degrees of freedom) reveal that both series appear to be non-stationary.[53] Moreover, the unemployment rate can be considered weakly exogenous from a statistical point of view (Table A4.2).

Table A4.2: Time Series Properties[54]

dgf	*dwage*	*ur*
	LR test for exclusion	
1	**5.02**	**7.03**
	LR test for stationarity	
1	**7.03**	**5.02**
	LR test for weak exogeneity	
1	**4.97**	0.72

Based on this, the analysis indicates no cointegration between the unemployment rate and wage inflation: the null hypothesis of $r = 0$, that is no cointegrating relationship, cannot be rejected at conventional levels (Table A4.3). Hence, the data cannot provide evidence of a long-run relationship between wage inflation and the unemployment rate.

[53]On theoretical grounds, the unemployment rate is bounded by the interval (0;1) and hence not truly I(1). The fact that it cannot grow out of bounds in the long run, however, does not preclude it from behaving like an integrated process in the shorter run, as evidenced by the test statistics. The stationarity tests presented above do not allow for a structural break in the series analyzed. The strong rise of the unemployment rate in the early 1990s—as described above—could be viewed as such a break. This proposition is not investigated further since a stationarity result for the unemployment rate when allowing for a break in the series even underscores the case for the NAWRU approach, see Schreiber and Wolters (2003).

[54]Bold test statistics indicate significance at the 5 percent level, the critical value with one degree of freedom being 3.84.

[55]L-max and Trace are the maximum eigenvalue and trace test statistics; L-max90 and Trace90,

Table A4.3: Cointegration Test Statistics[55]

Null	L-max	Trace	L-max90	Trace90
$r = 0$	8.67	10.23	10.60	13.31
$r \leq 1$	1.56	1.56	2.71	2.71

To confirm further the applicability of the NAWRU approach to the Finnish data, a number of additional considerations are of interest. The lack of cointegration between the two series could be due to a structural break in the cointegrating relationship during the observation period, in particular given the sharp rise in unemployment during the early 1990s. However, experimenting with various dummies did not soften the evidence against cointegration. Moreover, the limited number of observations used in the empirical assessment may introduce a small sample bias. Correcting for the bias, for instance along the lines of Cheung and Lai (1993), introduces even higher critical values, such that the hypothesis $r = 0$ would be accepted even more easily.

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instead, give the appropriate 90-percent critical values for r cointegrating vectors, see Johansen (1995), p. 215.

4.7.2 The Baseline NAWRU Model

The filtering process yielding the NAWRU is based on the unobserved components approach used by the European Commission, see Kuttner (1994), Denis *et al.* (2002), and Planas and Rossi (2003). In what follows, I outline the procedure for Finland as an exemplary case.[56] The starting point for the first component of the bivariate model is the definition:

$$U_t = C_t + \widetilde{T}_t, \tag{B1}$$

which decomposes the observable unemployment rate U_t in a cyclical component, C_t, and a non-cyclical, or trend component, $\widetilde{T}_t$. Additional exogenous regressors—$M \leq 3$—are assigned to the latter component, such that

$$\widetilde{T}_t = T_t + \sum_{m=1}^{M} a_m Z_{m,t}, \tag{B2}$$

where T_t represents the underlying long-term trend, or NAWRU. Without additional exogenous regressors, i.e., $M = 0$, the non-cyclical component and the NAWRU-trend coincide.

The trend component is modeled according to its statistical properties, i.e., no economic information (e.g., on structural breaks) is included. The most general specification (see further below) is given by a random walk (RW) with drift, where the drift term μ_t is itself a random walk (and the trend T_t, hence, a second-order random walk):

$$T_t = \mu_t + T_{t-1} + z_{1t}, \qquad with \tag{B3a}$$

$$\mu_t = \mu_{t-1} + z_{2t}. \tag{B3b}$$

[56]See the appendix of Billmeier (2004b) for more details, including the technical restrictions on the optimization technique.

Both errors, z_{it}, are n.i.i.d; if $Var(z_{1t}) = 0$, the model collapses to a first-order random walk with drift. On the other hand, the cyclical component in (B1) is specified as an AR(N) process:

$$C_t = \sum_{n=1}^{N} \phi_n C_{t-n} + \nu_t \tag{B4}$$

where $N \leq 2$, and ν_t is i.i.d. To guarantee stationarity of the cyclical component, it must hold that $\sum \phi_n < 1$.

The second component of the generic model is given by

$$\Delta\pi_t^w = \mu \underbrace{\left[+\sum_{l=1}^{L} \rho_l X_{l,t}\right]}_{a} \underbrace{\left[+\sum_{s=-1}^{S} \theta_s \Delta\pi_{t-s}^w\right]}_{b} \underbrace{[+\gamma(1-L)^p U_{t-1}]}_{c}$$

$$\underbrace{\left[+\sum_{r=0}^{R} \beta_r C_{t-r}\right]}_{d} + \underbrace{b_t}_{e}, \quad where \quad b_t = \sum_{i=0}^{I} \varepsilon_{t-i} \tag{B5}$$

This Phillips curve relationship links the change in wage inflation to (a) exogenous determinants of wage inflation, X_t, such as (changes in) labor productivity or (changes in) the terms of trade, with $0 \leq L \leq 10$ in the empirical application; (b) autoregressive terms of the wage inflation (with $0 \leq S \leq 2$); (c) the p-th difference of the lagged observed unemployment rate U_t; (d) the cyclical unemployment component C_t, (with $0 \leq R \leq 4$), and (e) an error term, which can have a MA(I) structure, $I \leq 3$.

In fitting the model to the Finnish data, the most generic model is used: with respect to equation (B1), the specification chosen is a bivariate autoregressive model, with the trend expressed as a second order random walk, hence, $Var(z_{1t}) \neq 0$ in equation (B3a). Furthermore, an AR(2) specification is selected for the cyclical component in (B4), as indicated by preliminary tests (not reported here). Experiments with additional exogenous regressors in (B2), i.e., $M > 0$, resulted in a deterioration of the statistical fit.

Regarding (B5), the second difference of the lagged first variable ($\Delta^2 U_{t-1}$), as well as the contemporary cyclical component of unemployment were included, that is, $d = 2$ and $R = 0$. The choice of a 2^{nd} order RW specification in (B3) implies for equation (B5) that $d = 2$, i.e., the lagged unemployment series regressor enters in second difference in order to obtain a stationary regressor. In (B5), no exogenous regressors were employed.[57]

In the Finnish case, the impact of different assumptions regarding the ARMA structure of the Phillips curve equation (B5) on the key parameters of the model is substantial. Only the ARMA(2,2) and (2,3) specifications yield a reasonable statistical description: normality assumptions on both equations are not violated at a 10 percent level. The estimated β_0 (in equation (B5)) indicates whether changes in wage inflation respond to the general economic environment as represented by the cyclical component of unemployment. This coefficient is significantly negative in both models, in accordance with the prior of a dampening effect of rising unemployment on the size of wage increases. Increasing the number of moving average terms in (B5) raises the t-value (in absolute terms). At 8.3 percent, the NAWRU derived for Finland in 2002 (in both specifications) is about 0.8 percentage points lower than observed unemployment.

For Finnish data, the final specification of the first component of the bivariate model (equations (B1), and (B3)) is, hence:

$$\begin{aligned} U_t &= T_t + C_t && \text{(B6a)} \\ T_t &= \mu_t + T_{t-1} + z_{1t} \quad with\ \mu_t = \mu_{t-1} + z_{2t}, Var\,(z_{2t}) \neq 0 && \text{(B6b)} \\ C_t &= \sum_{n=1}^{2} \varphi_n C_{t-n} + \nu_t && \text{(B6c)} \end{aligned}$$

[57]Maximization of the likelihood function was carried out by two algorithms. While the simulated annealing algorithm is slower than a Newton-type algorithm, it is more likely to identify a global maximum. Since our experiments showed that local maxima posed a problem using the Newton-type algorithm, the simulated annealing algorithm was applied throughout.

Using the ARMA(2,3) model, the final specification of (B5) is:

$$\Delta\pi_t^w = \mu + \sum_{s=-1}^{2} \theta_s \Delta\pi_{t-s}^w + \gamma\Delta^2 U_{t-1} + \beta_0 C_t + \sum_{i=0}^{3} \varepsilon_{t-i}. \tag{B7}$$

4.7.3 Tables

Table B4.1: Gap Signal Consistency

	HP100rt	PF	BQ	FD2-8
	Finland			
HP100rt	1			
PF	0.81	1		
BQ	0.46	0.51	1	
FD2-8	0.79	0.71	0.59	1
	France			
HP100rt	1			
PF	0.86	1		
BQ	0.62	0.57	1	
FD2-8	0.79	0.71	0.54	1
	Greece			
HP100rt	1			
PF	0.67	1		
BQ	0.46	0.46	1	
FD2-8	0.71	0.64	0.38	1
	Italy			
HP100rt	1			
PF	0.83	1		
BQ	0.73	0.65	1	
FD2-8	0.76	0.76	0.65	1
	United Kingdom			
HP100rt	1			
PF	0.81	1		
BQ	0.73	0.78	1	
FD2-8	0.74	0.62	0.68	1

Note: Statistic gives ratio of "same signs"; see section 4.5.2 for gap abbreviations.

Table B4.2: Forecast Performance for Alternative Gap Measures

	HP20rt	HP200rt	HP100	HP20	HP200	FD2-6	FD2-10
			Finland				
RMSE	1.97	1.94	1.94	2.12	1.57	1.37	1.36
Theil's U	1.29	1.26	1.27	1.38	1.03	0.89	0.89
DM	2.34	2.12	1.82	1.55	1.70	1.35	1.99
(p-value)	(0.99)	(0.98)	(0.97)	(0.94)	(0.96)	(0.91)	(0.98)
			France				
RMSE	0.60	0.74	0.85	0.72	0.87	0.54	0.52
Theil's U	0.87	1.07	1.23	1.04	1.26	0.78	0.75
DM	0.67	1.15	1.46	1.31	1.68	0.30	-0.12
(p-value)	(0.75)	(0.87)	(0.93)	(0.91)	(0.95)	(0.62)	(0.45)
			Greece				
RMSE	3.08	2.73	2.26	3.00	**2.17**	2.39	2.72
Theil's U	1.26	1.11	0.92	1.23	**0.89**	0.98	1.11
DM	1.77	1.04	0.04	0.87	-0.11	0.25	0.97
(p-value)	(0.96)	(0.85)	(0.52)	(0.81)	(0.46)	(0.60)	(0.83)
			Italy				
RMSE	1.07	1.24	1.29	1.25	1.24	**0.69**	1.11
Theil's U	0.95	1.09	1.14	1.11	1.10	**0.61**	0.97
DM	2.23	1.48	1.89	2.14	1.48	**-1.97**	1.21
(p-value)	(0.99)	(0.93)	(0.97)	(0.98)	(0.93)	**(0.02)**	(0.89)
			United Kingdom				
RMSE	**1.51**	**1.61**	1.84	2.14	1.87	1.86	1.97
Theil's U	**0.66**	**0.71**	0.81	0.94	0.82	0.82	0.86
DM	-0.64	-0.37	0.51	1.01	0.61	0.44	0.73
(p-value)	(0.26)	(0.36)	(0.70)	(0.84)	(0.73)	(0.67)	(0.77)

Note: Bold estimates indicate a (significantly) better performance than the univariate forecasting model; see section 4.5.3.1 for gap abbreviations.

Table B4.3: Evaluation of Forecast Performance (max. 1 lag)

	HP100rt	PF	BQ	FD2-8	NC	AR	UR
			Finland				
RMSE	1.91	1.87	1.19	1.48		1.14	2.62
Theil's U	1.24	1.22	0.78	0.97	1	0.74	1.71
DM	2.33	1.99	0.60	2.60			1.71
(p-value)	(0.99)	(0.98)	(0.72)	(0.99)			(0.96)
			France				
RMSE	0.61	0.92	0.55	**0.53**		0.53	0.66
Theil's U	0.88	1.34	0.80	**0.76**	1	0.77	0.96
DM	0.67	1.59	0.62	-0.33			0.96
(p-value)	(0.75)	(0.94)	(0.73)	(0.37)			(0.83)
			Greece				
RMSE	2.96	**1.78**	**1.58**	2.47		2.23	**1.94**
Theil's U	1.21	**0.73**	**0.67**	1.01	1	0.91	**0.79**
DM	2.06	-1.25	**-1.47**	0.85			-0.47
(p-value)	(0.98)	(0.11)	**(0.07)**	(0.80)			(0.32)
			Italy				
RMSE	0.97	1.00	**0.84**	**0.81**		0.90	0.91
Theil's U	0.86	0.88	**0.75**	**0.71**	1	0.80	0.80
DM	1.18	1.04	-0.72	-0.77			0.02
(p-value)	(0.88)	(0.85)	(0.24)	(0.22)			(0.51)
			United Kingdom				
RMSE	1.81	2.00	1.99	1.84		1.65	1.87
Theil's U	0.80	0.88	0.87	0.81	1	0.73	0.82
DM	0.57	1.12	1.19	0.63			0.79
(p-value)	(0.72)	(0.87)	(0.88)	(0.74)			(0.79)

Note: Bold estimates indicate a (significantly) better performance than the univariate forecasting model; see section 4.5.2 for gap abbreviations.

Table B4.4: Evaluation of Forecast Performance (max. 3 lags)

	HP100rt	PF	BQ	FD2-8	NC	AR	UR
			Finland				
RMSE	2.12	2.36	1.18	1.56		1.14	2.20
Theil's U	1.38	1.54	0.77	1.01	1	0.74	1.43
DM	2.54	2.33	0.49	2.67			2.35
(p-value)	(0.99)	(0.99)	(0.69)	(0.99)			(0.99)
			France				
RMSE	0.64	1.31	0.57	0.54		0.51	0.78
Theil's U	0.93	1.91	0.82	0.78	1	0.74	1.13
DM	0.85	3.04	0.58	0.82			1.86
(p-value)	(0.80)	(1.00)	(0.72)	(0.79)			(0.97)
			Greece				
RMSE	2.93	**1.65**	3.46	2.50		2.23	**2.13**
Theil's U	1.20	**0.67**	1.41	1.02	1	0.91	**0.87**
DM	1.92	-0.94	1.47	0.77			-0.14
(p-value)	(0.97)	(0.17)	(0.93)	(0.78)			(0.44)
			Italy				
RMSE	1.09	**0.86**	**0.95**	1.10		1.01	1.05
Theil's U	0.96	**0.76**	**0.84**	0.98	1	0.89	0.92
DM	1.23	**-1.32**	-0.95	0.54			0.18
(p-value)	(0.89)	**(0.09)**	(0.17)	(0.71)			(0.57)
			United Kingdom				
RMSE	**1.53**	2.07	1.99	1.84		1.73	1.88
Theil's U	**0.67**	0.91	0.87	0.81	1	0.76	0.83
DM	-0.56	1.15	0.94	0.39			0.59
(p-value)	(0.29)	(0.88)	(0.83)	(0.65)			(0.72)

Note: Bold estimates indicate a (significantly) better performance than the univariate forecasting model; see section 4.5.2 for gap abbreviations.

Table B4.5: Evaluation of Forecast Performance (max. 4 lags)

	HP100rt	PF	BQ	FD2-8	NC	AR	UR
			Finland				
RMSE	2.27	1.86	1.47	1.52		1.36	2.24
Theil's U	1.48	1.21	0.96	0.99	1	0.89	1.46
DM	2.14	1.72	1.15	1.31			2.14
(p-value)	(0.98)	(0.96)	(0.87)	(0.91)			(0.98)
			France				
RMSE	**0.64**	1.36	0.57	0.54		0.51	0.76
Theil's U	**0.93**	1.97	0.82	0.78	1	0.74	1.11
DM	0.85	2.99	0.56	0.82			1.72
(p-value)	(0.80)	(1.00)	(0.71)	(0.79)			(0.96)
			Greece				
RMSE	2.93	**2.21**	4.49	2.50		2.23	**2.13**
Theil's U	1.20	**0.90**	1.84	1.02	1	0.91	**0.87**
DM	1.92	-0.03	2.23	0.77			-0.14
(p-value)	(0.97)	(0.49)	(0.99)	(0.78)			(0.44)
			Italy				
RMSE	**0.97**	**0.96**	**0.95**	1.39		1.01	1.05
Theil's U	**0.86**	**0.85**	**0.84**	1.22	1	0.89	0.92
DM	-0.45	-0.56	-0.56	1.58			0.18
(p-value)	(0.33)	(0.29)	(0.29)	(0.94)			(0.57)
			United Kingdom				
RMSE	**1.53**	2.06	1.99	1.84		1.73	1.88
Theil's U	**0.67**	0.90	0.87	0.81	1	0.76	0.83
DM	-0.56	1.13	0.94	0.39			0.59
(p-value)	(0.29)	(0.87)	(0.83)	(0.65)			(0.72)

Note: Bold estimates indicate a (significantly) better performance than the univariate forecasting model; see section 4.5.2 for gap abbreviations.

Table B4.6: Evaluation of Forecast Performance (GDP deflator, max. 2 lags)

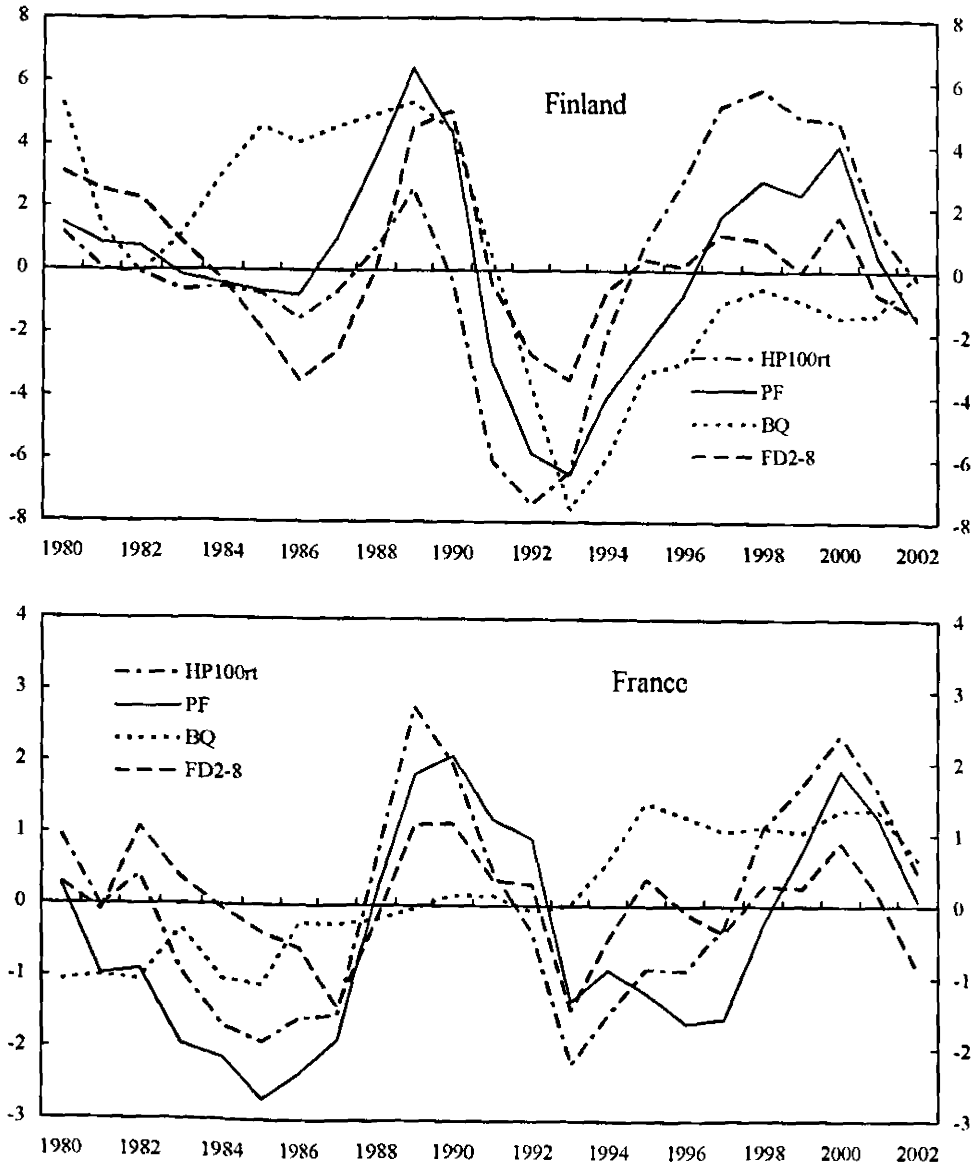
	HP100rt	PF	BQ	FD2-8	NC	AR	UR
				Finland			
RMSE	2.82	2.88	2.39	2.86		2.34	2.99
Theil's U	0.74	0.76	0.63	0.76	1	0.62	0.79
DM	1.08	1.93	0.04	1.58			2.66
(p-value)	(0.86)	(0.97)	(0.26)	(0.94)			(0.99)
				France			
RMSE	**0.45**	0.61	**0.49**	**0.48**		0.49	0.80
Theil's U	**0.70**	0.96	**0.76**	**0.75**	1	0.76	1.25
DM	-0.52	0.80	0.21	0.12			2.74
(p-value)	(0.30)	(0.79)	(0.58)	(0.52)			(0.99)
				Greece			
RMSE	4.44	**3.55**	3.66	3.84		3.56	6.31
Theil's U	1.44	**1.15**	1.19	1.24	1	1.15	2.05
DM	2.81	-0.03	0.15	1.56			3.86
(p-value)	(0.99)	(0.49)	(0.56)	(0.94)			(0.99)
				Italy			
RMSE	2.05	2.18	1.97	1.83		1.42	1.50
Theil's U	1.07	1.14	1.03	0.96	1	0.74	0.79
DM	2.45	2.68	1.56	3.25			0.29
(p-value)	(0.99)	(0.99)	(0.89)	(0.99)			(0.61)
				United Kingdom			
RMSE	1.33	1.36	1.43	**1.02**		1.21	**1.18**
Theil's U	1.18	1.21	1.27	**0.90**	1	1.08	**1.05**
DM	0.42	0.62	1.18	-0.61			-0.09
(p-value)	(0.66)	(0.72)	(0.88)	(0.27)			(0.46)

Note: Bold estimates indicate a (significantly) better performance than the univariate forecasting model; see section 4.5.2 for gap abbreviations.

4.7.4 Figures

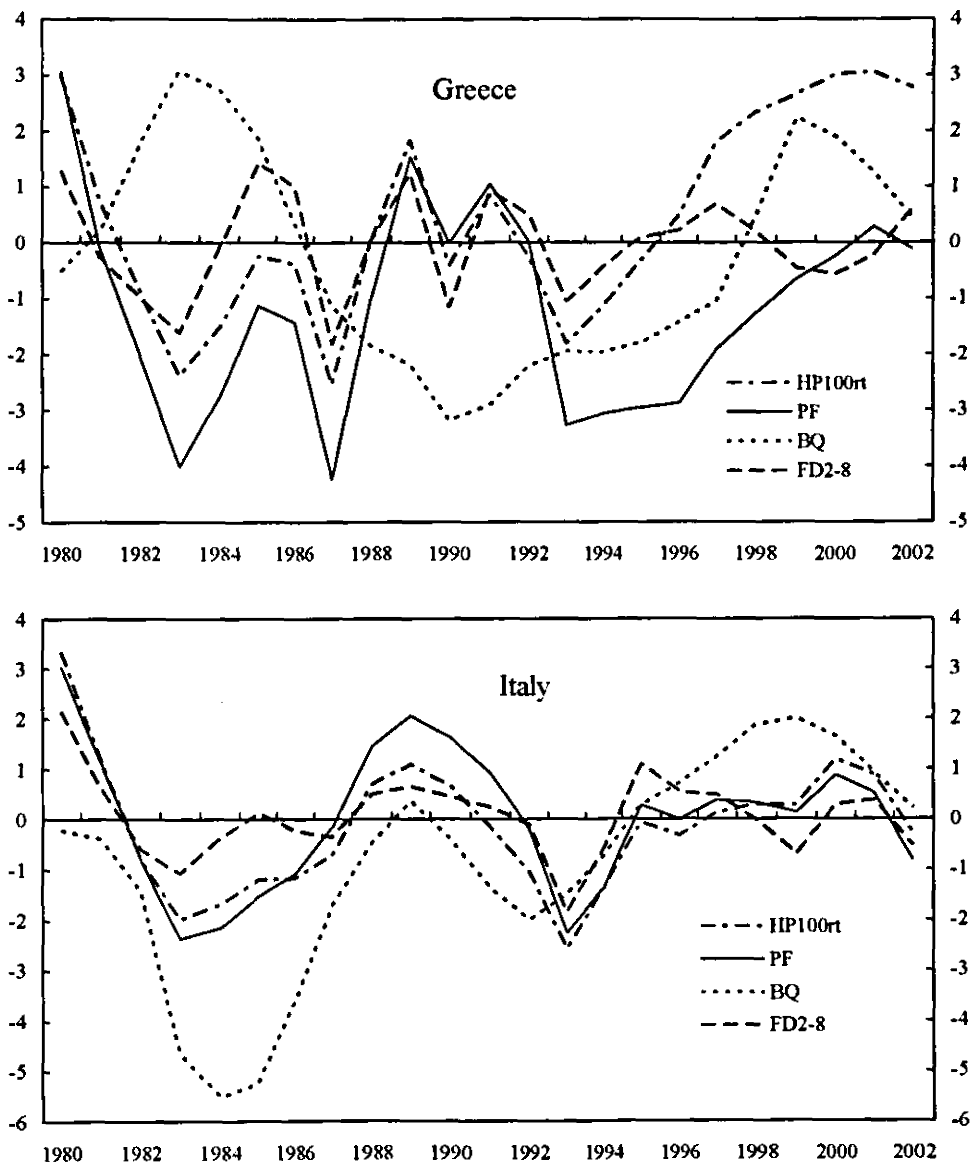
Figure A4.1: Output Gap Measures, 1980-2002

(In percent of potential)



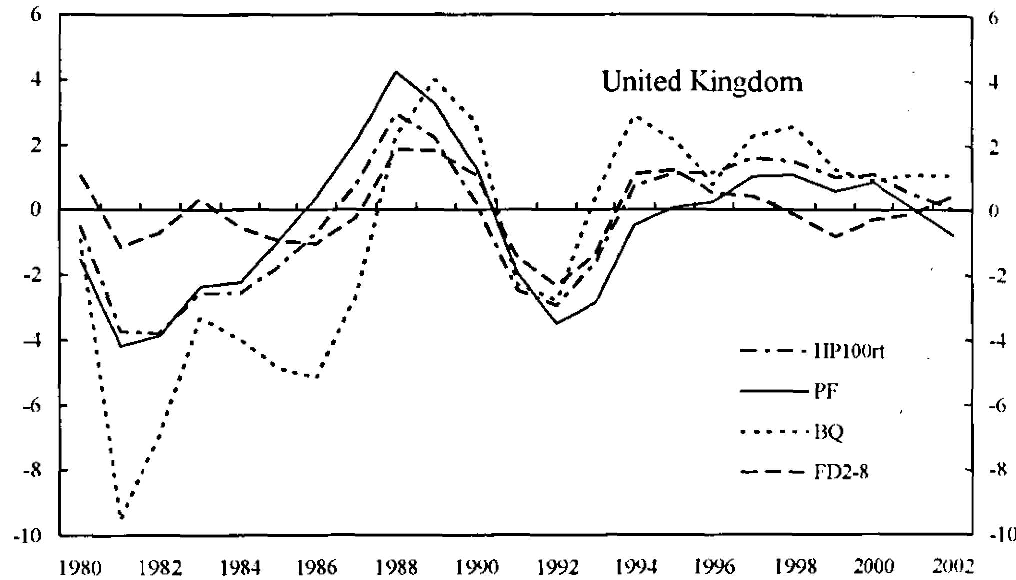
Sources: IMF *International Financial Statistics (IFS)*, European Comission; and author's calculations.

Figure A4.2: Output Gap Measures, 1980-2002
(In percent of potential)



Sources: IMF *IFS*, European Comission; and author's calculations.

Figure A4.3: Output Gap Measures, 1980-2002
(In percent of potential)



Sources: IMF *IFS*, European Comission; and author's calculations.

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