The interplay of public finances and the business cycle in Germany -

Evidence from time-varying VAR analyses *

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Abstract

In this paper we analyze empirically the timing of German fiscal policy over the business cycle based on quarterly data from 1970-2008. We take account of the endogeneity of the business cycle and public budget developments by employing vector autoregressions and allow for changing reaction patterns over time by applying time-varying parameter estimation techniques. Our analyses reveal four distinct regimes with important changes in the fiscal policy reaction. Overall fiscal policy (including automatic stabilizers and discretionary fiscal policy) turns out to have been strongly – but decreasingly – countercyclical in the first year after a shock, while the reaction with a longer forecast horizon was countercyclical in the 1970s and acyclical in later regimes. Our analyses furthermore indicate that the changes in the fiscal policy reaction have not been caused by a variation in economic volatility, but by structural changes in German public finances.

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1 Introduction: The interplay of fiscal policy and the cycle – the literature and our research approach

A central idea in standard Keynesian models is that a countercyclical fiscal policy helps to smooth the business cycle and to increase economic growth.[1] Hagen (1948) long ago stressed that the timing of fiscal policy is a crucial factor for these desirable effects and this is supported by more recent papers, such as those by Aghion et al. (2005) and Aghion and Marinescu (2008).

But what timing of fiscal policy do we really observe? Here, the findings of the empirical literature differ strongly. Aghion and Marinescu (2008), for example, find a countercyclical pattern of fiscal policy in the countries of the European Monetary Union (EMU), the UK and the US based on annual data from 1970 -2005.[2] Ballasone, Francese and Zotteri (2008), on the other hand, identify a procyclical pattern of fiscal policy for the EU 14 countries based on annual data from 1970-2004. This is supported by Ballassone and Francese (2004), who find a procyclical pattern of overall fiscal policy in the EU, the USA and Japan based on annual data from 1970-2000. Golinelli and Momigliano (2006) analyze overall fiscal policy in the EMU 11 from a real-time perspective based on annual data from 1988-2006, and find a countercyclical timing of fiscal policy. Studies like Gavin and Perotti (1997), on the other hand, argue on the basis of annual data from 1968-1995 that overall fiscal policy has been generally countercyclical in developed countries and more procyclical in developing countries.[3] This is supported by the study of Talvi and Vegh (2000), who argue that the strong procyclicality of fiscal policy in developing countries is caused by a higher volatility of the tax bases. For Ireland, Lane (1998) diagnoses procyclicality of Irish fiscal policy in a single-country time series study, while Muscatelli, Tirelli and Trecroci (2002) find for their complete sample of quarterly German data from 1971-1998 no significant fiscal policy response to shocks in output.

Instead of analyzing the overall reaction of fiscal policy, a large number of studies focus on one of the two sub-aggregates of overall fiscal policy: the "automatic reaction" of the budget deficit via the working of automatic stabilizers (for example, via the automatic increase in the tax burden caused by progressive taxation in times of upswings or by the automatic increase in unemployment transfers in times of downswings) and discretionary fiscal policy actions.

With respect to automatic stabilizers, there is a relatively strong agreement in the literature that automatic stabilizers are timed countercyclically in industrial countries. This is underlined, for instance, in recent papers by Debrun and Kapoor (2010) and Leigh and Stehn (2009).

With respect to discretionary fiscal policy, we find stronger controversies in the literature.[4] A serious problem in this respect is how to identify discretionary fiscal policy. The most common approach in the literature to identifying discretionary fis-

[1] Priesmeier and Stähler (2009) present a detailed survey of the literature discussing the effects of smoothing business cycles on economic growth.

[2] See as well the discussion in von Hagen et al. (2001).

[3] Alesina, Tabellini and Campante (2008) offer a rationale for this finding: developing countries have larger problems with political corruption and the larger the problems with corruption, the stronger are the incentives for voters to reduce political rents by demanding more public goods or lower taxes in booms. Along similar lines, Calderón et al. (2004) show that developing countries with better institutions are more able to conduct fiscal policy countercyclically. Another reference showing procyclical fiscal policy in Latin America is Stein et al. (1999).

[4] Golinelli and Momigliano (2009) compare explicitly the effect that cyclical adjustment of the primary balance has on the results of studies on the timing of fiscal policy.

cal policies is to “cyclically adjust” revenue and expenditure developments (applied, for example, in Alesina and Perotti (1995) and Giavazzi and Pagano (1996)).[5] The findings in this literature differ strongly. Gali and Perotti (2003) found that cyclically adjusted deficits have not reacted to the business cycle in Europe after the signing of the Maastricht treaty. These results are at least partly confirmed in studies such as Ballabriga and Martinez-Mongay (2002) and Wyplosz (2006). Other studies, like Deroose, Larch and Schaechter (2008), compare the timing of fiscal policy in the euro area and the US since the mid-1990s and find that discretionary fiscal policy in the euro area has tended to be more procyclical. Separate analyses of cyclically adjusted revenue and expenditure developments seem to indicate that, although the fiscal balance might point at acyclicality of fiscal policy, discretionary revenue policies seem to be timed countercyclically, while discretionary expenditure policies tend to have been procyclical in Europe (see, for example, Fatás and Mihov (2009) or Turrini (2008)).[6]
The widely varying findings in the literature call for a more in-depth analysis of the timing of fiscal policy. From our perspective, we perceive mainly three problems with the discussed studies:
The first problem is that the existing studies only scarcely apply time series analyses – which would seem necessary to account for the dynamic interdependencies between the business cycle and the public budget.
The second problem is that even those studies that do use time series data, focus on annual data (Muscatelli, Tirelli and Trecroci (2002) are an exception). This not only strongly restricts the number of observations but also complicates the accurate separation of shocks and their effects. An output gap shock does not need a whole year to affect the economy and fiscal policy reacts far more quickly than at an annual frequency. Based on yearly aggregates, the short-term reaction cannot be analysed and the analysts often need to confine themselves to static econometric approaches that do not allow for any intertemporal dynamics and interpretations. This speaks strongly in favour of higher frequency - in other words, quarterly - data.
Third, the interplay of fiscal policy with the economy is likely to be subject to change over time. Empirical support for this proposition comes, for instance, from the study of Aghion and Marinescu (2008), who find that the countercyclicality of the overall fiscal policy response increased in the US from 1960-2005 and in the UK from 1970-2005, but significantly decreased in the EMU countries.[7] Additionally the results for different subsamples in the study by Muscatelli, Tirelli and Trecroci (2002) – who find

[5]Cyclical adjustment usually means subtracting a ”cyclical” component from the aggregate revenue and expenditure developments. The cyclical component is calculated based on a state-indicator of the business cycle (usually the output gap) and an elasticity measure for the effects of business cycle fluctuations on fiscal developments (see for a calculation of such elasticity measures, say, Girouad/André (2005) or Mohr et al. (2001)).

[6]Some authors argue that cyclical adjustment methods are unable to unmask discretionary fiscal policy (see, for example, Chalk (2002) or Larch and Salto (2005)). One argument in this respect is that the elasticity of fiscal variables to the macroeconomic development might not be invariant over time (see, for example, Jaeger and Schuknecht (2004)) or might not cover all the relevant cyclically sensitive spending (or revenue) categories (see, for example, Darby and Mélitz (2008)). One further aspect that is stressed by some authors is that the timing of fiscal policy does not only need to be analyzed ex post, but also has to be based on the information available at the time when the measures are adopted. This literature relies on real-time data on budget plans and expectations about the macroeconomic development. The number of studies in this category is still relatively limited, but so far the results tend to indicate that cyclically unadjusted fiscal plans show a stronger countercyclicality than ex-post data (see, for example, Beetsma and Giuliodori (2008), Giuliodori and Beetsma (2007) , Cimadomo (2007) or Golinelli and Momigliano (2009) for related studies).

[7]Aghion and Marinescu (2008) use a static state space model based on annual data, which delivers information only on the time variation of the contemporaneous relationship.

no significant overall fiscal policy response to shocks in the output gap from 1982-1998, but a significant countercyclical reaction in the subsample from 1971-1982 – point to a response pattern in German public finances that has changed over time.[8] It might therefore be the case, for example, that institutional reforms or changes in the general orientation of fiscal policy under different governments change the way public finances react to economic fluctuations (see, for instance, Duncan and Schmidt-Hebbel (2004)). On the other hand, it might be that the economic structure changes: economic shocks could lead to stronger output fluctuations which, in turn, trigger stronger fiscal policy reactions, or output shocks themselves might have decreased or increased in size. As Talvi and Vegh (2000) demonstrated, this stronger volatility in output and – therefore – in the tax bases might have important effects on the countercyclicality of the budget. However, the existing studies are barely able to address such time-variation in the public budget's reaction to innovations in the output gap.

In this paper, we want to contribute to the empirical analysis of the interplay of public finances and the business cycle. We focus on aggregate data for budget and business cycle developments to cover the whole spectrum of fiscal policy – automatic stabilization as well as discretionary fiscal policy. This has important advantages and disadvantages. While the overall perspective does not allow us to distinguish between automatic stabilizers and discretionary policy, it avoids biased results which could result from an inaccurate method of cyclical adjustment. Given the lacking research on the characteristics of automatic stabilizers for quarterly data, the danger of applying an inaccurate method of cyclical adjustment is large. Furthermore, we rely on a large time-span of data and aim at a time-varying analysis of fiscal policy, for which we would need a time-varying method of cyclical adjustment that is well able to distinguish between variation over time that comes from the automatic stabilizers and variation of discretionary fiscal policy.[9]

Our approach tackles the general endogeneity problem, and accounts for the fact that fiscal policy develops over time by applying appropriately just-identified vector autoregressive (VAR) methods. To account for possible regime changes in the timing and the size of fiscal policy responses in different periods, we include a regime-switching model as well as time-varying VAR analyses that permit the coefficient matrices to evolve over time (as random walks). The data builds on the Deutsche Bundesbank national accounts database in quarterly frequency from 1970-2008, which gives us a rather long and high-frequency dataset suitable for reliable time series appliciations. Taken together, we wish to contribute to answering the following three questions from a dynamic perspective on quarterly data:

1. Do we observe a stable cyclical reaction pattern or has it changed over time and – if so – how and when?

2. If there are systematic changes in the fiscal policy reaction, have they evolved gradually over time or occurred abruptly?

3. What does the data tell us about the reasons for possible changes? Do they stem from changes in the economic volatility or changes in the structure of German public finances?

[8]Based on a time-varying VAR, the authors cannot find any evidence for changing fiscal policy regimes after 1980. They cannot evaluate the period before 1980 due to a long burning in period in their estimation procedure.

[9]See footnote 5.

The paper proceeds as follows: In part two, we briefly review different concepts of the cyclicality of fiscal policy and state the reasons for our definition of cyclicality. Part three presents the data and the indicators employed. In part four, we study the timing of fiscal policy based on a time-invariant benchmark approach, which we find not to be the most appropriate choice as there are important indicators pointing to parameter instability even in augmented models that allow for discrete regime changes. To account for this finding, we perform time-variant analyses in part five, which point to different regimes of fiscal policy timing that gradually flow into each other. The last part concludes.

2 Concepts of the cyclicality of fiscal policy

Generally, fiscal policy can be unrelated with the cycle (acyclical) or respond in a pro- or countercyclical way to economic developments. To assess the relationship of fiscal policy and the business cycle, a measure for the business cycle, a measure for the fiscal policy stance and finally a definition for acyclical, counter- and procyclical fiscal policies is needed. In the literature, we can find different concepts for categorizing the timing of fiscal policy.

With respect to the variables, the broad majority of studies use the output gap as an indicator for the business cycle (see Golinelli and Momigliano (2009, p. 42 ff.) or OECD (2010, p. 9 ff.))[10], as it divides economic development into phases in which output is below potential output (output gap smaller than zero), phases in which output is above potential output (output gap larger than zero), and phases in which the output equals potential output. This gives a clear-cut picture of “good” and “bad” economic times, while other measures – such as GDP growth rates – make additional and arbitrary definitions necessary.

Fiscal policy is usually measured by the primary balance. Here, we can distinguish between studies, which try to assess overall fiscal policy (Balassone, Francese and Zotteri (2008), Golinelli and Momigliano (2006) and Ballabriga and Martinez-Mongay (2002)) and those, which try to separate discretionary fiscal policy by cyclically adjusting the primary balance (see, for example, Turrini (2008) or Golinelli and Momigliano (2009)). As we want to analyze overall fiscal policy (see discussion in part I), we chose the unadjusted primary balance as well-suited indicator for fiscal policy.

Concerning the relationship of the economic cycle and fiscal policy, the literature provides three different approaches to defining cyclicality:

i) Related to standard Keynesian approaches as well as the theory of tax smoothing (see Barro (1979)), it seems straightforward to use the output gap and the primary balance in levels - as, for example, Muscatelli, Tirelli and Trecroci (2002) and Aghion and Marinescu (2008) do. In this case, a positive output gap reflects a boom. When the primary balance is positive (equalling a fiscal surplus) at the time of a positive output gap, the fiscal stance can be expected to contribute to smoothing the business cycle and can therefore be called “countercyclical”. For the case of a negative output gap, a countercylical fiscal policy would demand negative primary balances. Fiscal policy would be termed “acyclical” if the primary balance is zero despite an output gap different from zero or if the output gap is zero but the primary balance differs from zero.

ii) Alesina and Tabellini (2005) and Turrini (2008), for example, propose relying on the change of the primary balance instead of the primary balance itself. This results

[10] Some studies use instead the growth rate of output as indicator. See, for example, Lane (2003).

in a different definition of pro- and countercyclicality. In this case, countercyclicality means that the primary balance increases while the output gap is positive and decreases while it is negative (vice versa for procyclical fiscal policy). Acyclical fiscal policy would imply that the primary balance changes, although the output gap remains at zero, or that the output gap is different from zero, but the primary balance remains unaltered. In some cases, this definition is weaker than the first, as a negative primary balance can be classified as countercyclical under a positive output gap – if only the balance improves. This implies a less clear link between countercyclical fiscal policy and the smoothing of the cycle.

iii) A third approach is to employ changes of both the output gap and the primary balance as variables (see, for example, Lane (2003) or Leigh and Stehn (2009)).[11].In this case, countercyclical fiscal policy is defined as a positive (negative) change in the fiscal balance in case of a positive (negative) change in the output gap, while procyclical fiscal policy would be given if the fiscal balance worsens (improves) when the economic situation improves (worsens). Acyclical fiscal policy would imply that the primary balance changes although the output gap remains constant or that the output gap changes but the primary balance remains unaltered. The main problems of this third approach are that the link between fiscal policy and the business cycle is not clear cut and – as it abstracts from the level of the variables – a lot of information is lost. Situations in which the primary balance is negative and the output gap is positive would, for instance, be classified as countercyclical, as long as the primary balance and the output gap both show positive changes.

In this paper, we apply the first approach. In our view it is the strictest and most reliable approache as it establishes a straightforward, clear and theory-based link between the state of the business cycle and fiscal policy.

3 Measurement concepts and data-set

According to our definition of cyclicality, we need the output gap as indicator of the state of the business cycle and the overall primary balance as indicator for the fiscal policy stance.

For the output gap variable, we have calculated the real GDP gap based on the quarterly national accounts database of the Deutsche Bundesbank. Nominal GDP was first realized by the chain-linked GDP deflator and then seasonally adjusted.[12] In a second step, we applied the Hodrick-Prescott Filter (Lambda=1600) to the real GDP series which we prolonged with its own linear trend in the past (1960-1970) and the future (2009-2019) to avoid a distortion of the results at the lower and upper bounds of our series.[13] The real output gap variable was then calculated as the difference between actual real GDP and potential real GDP (measured by the HP filtered trend) as a percentage of potential GDP.[14]

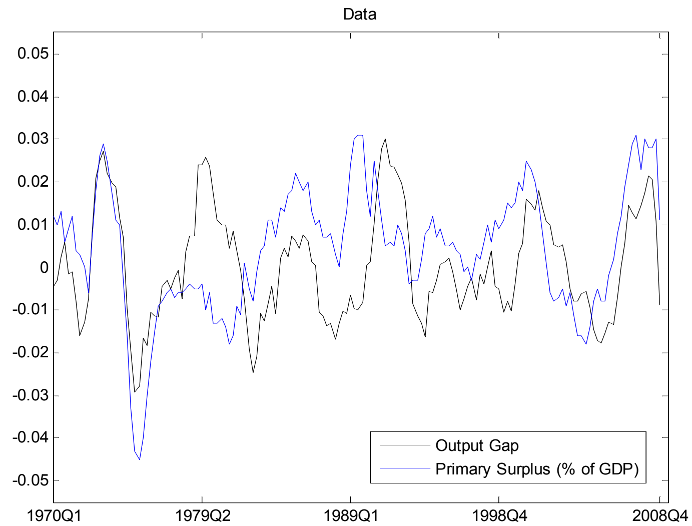
[11] As the change of the output gap is closely related to the growth rate of output (the growth rate of output is basically the change in the output gap in percentage points plus the trend growth rate), some studies use the output growth rate instead of the change of the output gap (see, for example, Formi and Mormigliano (2004)).

[12] We used the BV 4.1 procedure of the Federal Statistical Office to adjust the series for seasonal effects.

[13] We also applied an one-sided Hodrick-Prescott Filter, which did not affect our results substantially.

[14] This measure corresponds to the difference between log actual and log potential output, for example used by Muscatelli et al. (2002).

Figure 1: Fiscal Policy over the business cycle in Germany 1970-2008



With respect to fiscal policy we want to distinguish between expansionary and restrictive fiscal policies based on the overall primary balance. Here, we decided to focus on general government including social security insurances because revenue and expenditure developments in the social security system are subject to political discretion as well and tend to affect the overall fiscal stance of the government, which, in turn, influences macroeconomic development. However, we have excluded interest spending because the government's ability to change this spending category is very limited and interest should therefore not be included in the discretionary fiscal policy reaction. Moreover, it should be noted here that the general timing pattern of fiscal policy over the business cycle is – based on quarterly data – only slightly affected by interest spending, which indicates that we obtain very similar results even if we do not subtract interest spending.

The exclusion of interest spending from the fiscal policy reaction is consistent with other approaches in the literature (see, for example, Perotti (2004)), but not uncontested as Blanchard and Perotti (2002), for instance, include an interest component. Other studies, which do not abstract from interest spending, often include an additional debt feedback variable (see, for example, the discussion in Golinelli and Momigliano (2009)). For our series of the real fiscal stance of the government, we have subtracted general government expenditures (excluding interest) according to the national accounts definition from general government revenues (mostly taxes and social security contributions). Thus, our fiscal stance is a primary surplus variable. To derive a measure of the real and seasonal adjusted fiscal stance we have first realized the expenditure and the revenue series with the chain-linked GDP deflator and then adjusted the series seasonally. In a second step, we derived the primary surplus ratio as a percentage of real trend GDP. Figure 1 shows the development of the business cycle (real GDP gap – black line) and the government fiscal stance as the primary balance ratio (blue line). We see already that the two stationary series seem to move very closely together in some periods, while the series diverge strongly during others. This can be seen as a first indication that a time-invariant analysis might not be the optimal choice for this data.

4 Time-invariant VAR analyses: A benchmark model

Our aim is to analyze the timing of fiscal policy over the business cycle – an attempt which is directly affected by the endogeneity of fiscal policy and the business cycle. To tackle this problem in an intuitive way and to account for the fact that fiscal policy develops over time, we propose time series analyses based on vector-autoregressions (VAR). We start our investigation with a time-invariant benchmark model.

4.1 Identification of an output gap shock

To analyse the structural relations between the business cycle and fiscal policy we assess the impact of identified exogenous and unanticipated shocks because these do not affect the systematic relations between the aggregates.[15] There are essential problems of such a strategy. We have to make sure that the reactions reflected, are, in fact, due only to the considered economic shock, which is usually given if the disturbances of the estimated system are instantaneously uncorrelated and thus the residual variance-covariance matrix of the estimated process is diagonal. However, more than one parametrization is possible to generate uncorellated residuals within a just-identification scheme and each parametrization identifies a different structure of the underlying economy.[16]

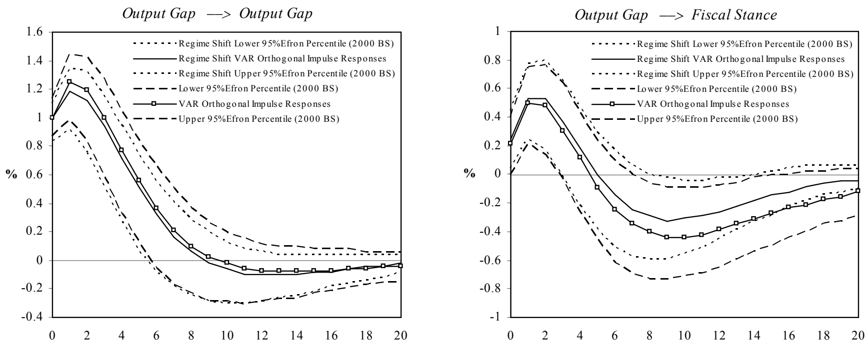
To keep our scheme simple, we apply the commonly used assumption that identification of the system can be achieved by implementing an economic structure only on the contemporaneous interactions of the reduced form residuals and not on the contemporaneous relations of the variables themselves, which corresponds to the B-SVAR approaches (Bernanke (1986)). In a next step, we use the Cholesky-type variance-covariance decomposition introduced by Sims (1980) to orthogonalize the residuals.[17] This approach implies a recursive structure of the economy (represented by the recursive structure of the residuals). In our case, this structure is well supported by additional information on the timing of the German tax and transfer system and on political decision lags in representative democracies. Given the variable ordering the fiscal stance reacts contemporaneously to the output shock, whereas there is no feedback reaction running from the structural primary balance to the output gap within the same quarter. This modelling of a contemporaneous reaction of public finances on changes in the output gap is commonly applied in SVAR studies on fiscal policy issues. According to the seminal work of Blanchard and Perotti (2002), the instantaneous interactions can be completely assigned to the working of automatic stabilizers, because reactions in discretionary policy do not occur in the same quarter

[15] Structural shocks and the identification scheme are assumed to be time-invariant over the sample. With respect to the estimated residuals of the constant parameter VAR and measurement residuals of the time-varying parameter model, this seems to be a reasonable specification because the structure of the residuals does not change over time.

[16] Ideally, the identification is based on economic theory or schemes that are well-established in the literature.

[17] In general, empirical studies that consider fiscal policy issues refer to one or more of four main identification approaches (see Caldara and Kamps (2008)). First, the standard recursive approach introduced by Sims (1980) and applied in the context of fiscal policy analysis by Fatás and Mihov (2001). In fact, this approach is most frequently used in fiscal and monetary policy applications of VARs and TVP VARs; see, for example, Muscatelli et al. (2002, 2007), Baumeister, Durinck and Peersman (2008) or Koop and Korobilis (2009). Second, the structural VAR approach by Blanchard and Perotti (2002) and Perotti (2005, 2007); third, the sign-restriction method by Uhlig (2005); and, fourth, the event-study method by Ramey and Shapiro (1998). Alternatively, there exist approaches that distinguish between short- and long-run structural shocks (see, for example, Blanchard and Quah (1989) or Lee and Chin (2006)).

Figure 2: Benchmark model impulse responses



of the shock owing to political decision-lags in representative democracies like Germany, where the parliament needs to be involved in the decision-making process. On the other hand, the instantaneous zero-feedback reaction implied by the scheme is not beyond all doubt. However, owing to the large number of studies (see, for example, Muscatelli et al. (2002) as a reference) that consider the contemporaneous non-zero reaction of primary surpluses to be more important, it seems to be reasonable for us to rely on this restriction in our recursive identification scheme.[18] The decomposition leads to a highly significant contemporaneous reaction of the fiscal policy aggregate to a 1 percentage point innovation in the output gap.[19] This reaction, which we assign completely to automatic stabilization, equals a value of 0.215 (0.077 (percentage points 0.245 (0.082) percentage points for the regime-switching model).[20]

4.2 Benchmark structural analyses results

We used multivariate least-squares (LS) estimation to obtain time-invariant values for the coefficients of a two-dimensional VAR(2) with constant. The optimal lag order was set according to information criteria, autocorrelation analysis and with respect to the fact that the time-invariant VARs will be taken as reference for the time-varying parameter models that may be overfitted by implementing high lag-orders. To analyse the impact of the identified output gap shock on the dynamics of fiscal policy over longer horizons, we compute impulse response functions that can be interpreted as forward-looking budgetary reaction functions over a horizon of 20 periods. These show how the primary deficit reacts to an output shock until the system gets back into its stable equilibrium. Figure 2 (continuous line with quads) shows the corresponding impulse-response functions of the output gap itself and of the overall fiscal stance on a positive 1 percentage point shock in the output gap within 95% Efron confidence intervals based on 2,000 bootstrap iterations (dashed lines). The left panel shows the reaction of the output gap to a shock in itself. Here, the one-off impulse leads

[18]A different ordering of the variables did not change the overall responses of the surplus and the output gap, but the contemporaneous reactions.

[19]Standard-deviations of the reactions are estimated by maximum likelihood using the scoring algorithm of Amisano and Giannini (1997) for a general B -model, in which the no-contemporaneous-feedback restriction is exogenously set.

[20]Isolation of the fiscal components is contemporaneously achieved. For the following quarters, the decomposition of systematical automatic stabilization and systematical discretionary policy is much more complicated, because a sophisticated decomposition approach is needed. Leigh and Stehn (2009) offer a political decision lag approach. For longer horizons Du Plessis and Boshoff (2007) apply the cyclical adjustment procedure of Girouard and André (2005) on quarterly data for South Africa.

to a short positive upward dynamic with higher GDP growth rates than trend GDP growth rates (which leads to a further widening of the output gap) before the growth rate of actual GDP starts to fall below the trend growth rate between period one and two and the positive gap thus slowly begins to close. After six quarters the output gap ratio is no longer significantly different from zero. The point estimate decreases further until the reaction becomes slightly but insignificantly negative from quarter ten on, where actual output falls below trend output.[21] Afterwards, the system tends to the equilibrium where actual output equals trend output again.

The primary balance ratio (right panel) reacts immediately and significantly positively to the output gap shock with a contemporaneous increase of 0.22 percentage points (which can be interpreted as the isolated working of automatic stabilizers, as we have assumed political decision lags of at least one quarter). Thereafter, the surplus ratios increase further to the maximum of around 0.50 percentage point between the first and the second quarters. After the second quarter, the surplus ratio decreases and is not significantly different from zero between quarters three and four. Around quarter seven, the surplus ratio even starts to turn into a deficit, which reaches its highest value with -0.44 percentage points in period nine. From then onwards deficit ratios start to fade out – parallel to the adjustment of output to the new equilibrium level.

If we bring together the development of the output gap and of the fiscal stance to evaluate the timing of fiscal policy we see – according to our definition (compare part 2) – a clearly countercyclical reaction of fiscal policy in the first three quarters as significant surpluses are accompanied by positive output gaps.[22] Thereafter, the faster decrease of surpluses leads to a time span (quarters five to nine) where we observe insignificant and, from quarter six on, significant deficits, although the output gap is no longer significantly different from zero, which would indicate acyclical fiscal policy. Taken together, the benchmark model indicates a fiscal policy, which is first strongly countercyclical and then – after the second year after a shock – acyclical.

The feedback effects of the fiscal stance on the output gap (not shown in the figures above) are clearly insignificant and rather non-Keynesian, as the impact of increasing surpluses on the output gap seems to be positive. However, this finding is well in line with existing VAR studies for the German case (see, for example, Muscatelli et al.(2002)).

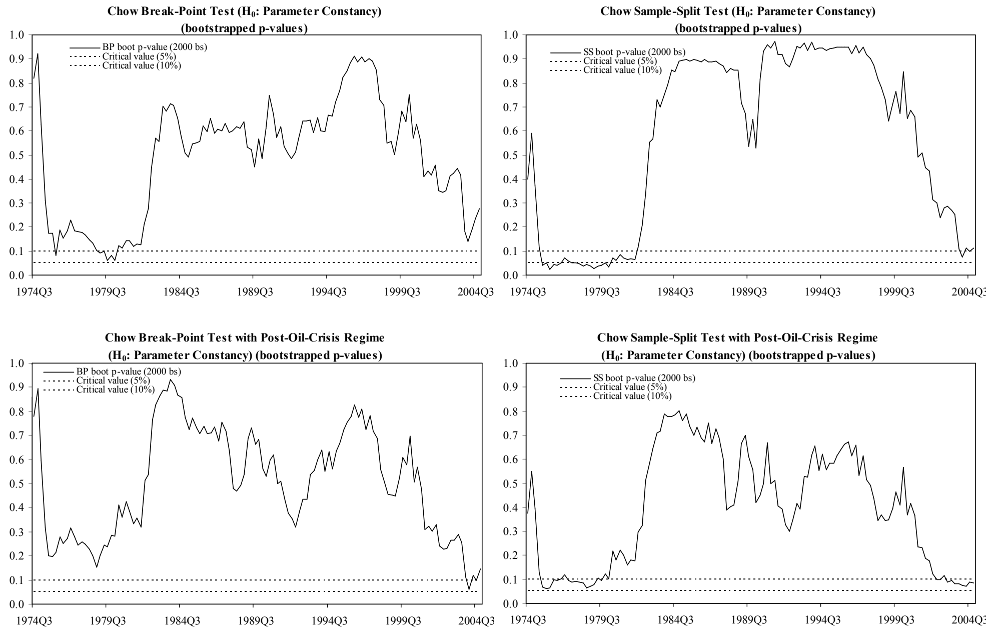
4.3 Parameter stability: Constancy, discrete switches and gradually evolving fiscal policy regimes

As we consider a rather long sample horizon from 1970 to 2008, we expect some variation in the reactions to output gap shocks. For instance, variations can result from political and institutional regime changes or, alternatively, from changes in the structure of the economy and the way in which it reacts to economic shocks. Basically, we can make a distinction between two different types of parameter changes. On the one hand, regimes can switch abruptly and time point-specifically (due, for example, to abrupt and far-reaching structural economic or political changes). On the other hand, changes in the fiscal structure can occur more gradually (for example, if the timing of fiscal policy is adjusted in small steps or economic agents adjust only

[21] It is not uncommon for the economy to go through such a period of underutilization before the new equilibrium level is reached. This might result from expectation or inventory adjustment effects.

[22] If we take the arguments of political decision lags (see e.g. Leigh and Stehn (2009)) and the findings from VARs of higher dimension (presented in appendix A), which additionally include variables describing possible transmission channels of the output gap shock, into account, this countercyclical reaction is likely to be dominated by the working of countercyclical automatic stabilizers.

Figure 3: Break-point and sample-split tests



slowly). Each type requires a different modelling. To be able to identify the properties of regime changes that are potentially included in our time series, we test for parameter constancy using a wide range of specifications. In a first step, we test for parameter constancy applying conventional Chow-type tests on the estimated time-invariant VAR.[23] Using bootstrap versions of the break-point and the sample-split test (2,000 iterations) enables us to identify possible periods of structural breaks. The bootstrapped p-values for every quarter of our dataset from 1974:3 to 2004:4 are plotted in figure 3.[24]

In the first row of figure 3, the results without any additional deterministic terms are presented. Both tests indicate parameter instability at the beginning of the sample – the sample-split test that excludes the estimated residual variance, even on a 5 per cent level. In the second step, we analyse whether this instability results from abrupt regime switches, which can be identified exogenously. The Chow-type tests indicated instability within the period from 1975 to 1982, during which the second oil-crisis led to massive economic distortions. Therefore, in the first instance we implement impulse dummies in 1979:2 and 1982:3 to capture the transitory effects of the oil price shocks on the output gap and the fiscal stance.[25] Then we checked whether there is evidence

[23]Break-point test (BP) and sample-split test (SS), compare the residual variance estimate from a model with constant parameters with the residual variance from a model that allows for a change in the parameters at a specific point of time. The break-point test has the null hypothesis that the VAR coefficients, the deterministic terms and the residual variance do not change. The sample-split test has the null hypothesis that only the coefficients and the deterministic are constant. According to Candelon and Lütkepohl (2001) the corresponding test statistics perform very poorly in smaller samples. We follow the authors suggestion to use bootstrap versions of the tests.

[24]Computations at the beginning and the end of the sample have to be interpreted very cautiously because there are only a few observations available for the corresponding sub-sample estimations. Testing is based on a minimum of 16 observations in the sub-samples.

[25]By implementation of these impulses we are able to capture the strong boom just before the first oil-price shock, which pushed Germany's output gap up, as well as the end of the strong recession that followed this boom. Both dummies are found to be highly significant on a 5 per cent level but only in the output gap equation. The first impulse estimate is 0.019 (3.733), the second -0.013 (-2.452).

for significant regime switches following the second oil crisis and tested the hypothesis of a partisan regime shift in 1982:4 when the conservative-liberal government came into power and lasted until 1998:4. Furthermore, we tested the hypothesis of a post-oil-crisis regime starting in 1982:4 in two different versions: the first assuming one regime for the remaining sample horizon, the second assuming one regime only until the fall of the Berlin Wall (1989:4). On a 5 per cent level none of the regimes can be significantly identified. On a 10 per cent level, the post-oil-crisis regime which lasts until reunification is estimated to have marginal smaller output gaps (-0.002). Overall, the implementation of the identified regime shift and the additional oil crisis deterministic leads to a slightly better fit of each equation, according to the adjusted coefficients of determination, 0.86 (0.84) and 0.90 (0.89).

Based on the same identification scheme as in the benchmark case, figure 2 shows the impulse-response functions for the regime-switching VAR (continuous line and dotted lines), which indicate that the implementation of the identified deterministic leads to a slightly stronger countercyclical fiscal policy reaction to a 1 percentage point output gap shock in the beginning and less strong and significant cyclical reactions later on. The structure of the impulse-response, however, does not change.

After implementation of the "oil crisis to reunification regime" and the transitory oil-crisis dynamics into the model, the corresponding break-point tests shown in the second row of figure 3 now reject the null hypothesis of instable parameters, whereas the sample split tests still indicate some instability at the beginning of the sample.[26] Thus, the regime-switching model that allows only for discrete parameter switches may not capture the full change in the data-generating process and may therefore lead to incorrect conclusions concerning the structural relations of the economic variables. Changes may occur more gradually than is supposed under this modelling. In fact, the recursive parameter estimates in their confidence intervals – shown in appendix B for the other variable in each equation of the regime-switching model – have different levels at the beginning of the sample than at the end. In particular, within the first decade (1970 to 1980) and around the time of German reunification in 1990/1991 there is some significant more or less gradual change to observe. Thus, a model that covers one or more gradual regime changes seems to be more appropriate for the data. In the next section, we will adopt an even more general model that allows for such gradual evolvements of parameters during the sample period, i.e. what is known as a time-varying parameter (TVP) VAR.

5 A time-varying parameter (TVP) VAR in a state space framework

The evidence for parameter instability in time-invariant models speaks in favour of a time-variant model, where the coefficients and thus the corresponding impulse responses can differ gradually over time. Other approaches based on the implementation of discrete structural breaks (see, for example, Stock and Watson (1996)) have been rejected. Forming sub-samples for each fiscal policy or business cycle regime is a possi-

Alternatively, we implemented different regimes for the time between the two oil crises using shifts that reach from 1975-1982, 1977-1980, 1979-1982 and 1980-1982. None of them is significant.

[26] In addition, we tested for parameter stability within the identified sub-samples, 1970:3 to 1982:3 and 1982:4 to 2008:4. Within the sub-samples there are signs of parameter instability as well – in particular in the longer sample – according to the Chow-type tests. However, such sub-sample VARs would be the most inappropriate choice because the problem of potentially neglected gradual changes remains valid and is even combined with a significant loss of observations.

ble alternative, but would lead to very short samples of only a few observations, which significantly reduces the reliability of the estimates. The most frequently used method for incorporating time variation is to employ a time-varying parameter (TVP) VARs (see, for example, Doan, Sims and Litterman (1984), Primiceri (2005) or Muscatelli, Spinelli and Trecroci, (2007)).[27]

5.1 Normal linear state-space representation

In situations of time-varying relations between economic variables, simplification can be achieved by state-space frameworks, which allow the inclusion of additional information on the stochastic behaviour of the reduced-form VAR coefficients over time. The main idea is that the VAR coefficients – from now on called "states" – can be calculated recursively from measurable data described by a certain data-generating process, assuming that the states follow a stochastic process with *a priori* known properties.

In other words, the state-space model is a two-layer model, where the external layer involves the measurable data described in a measurement equation, and the internal layer involves the additional information on the path of the states in form of a state equation. According to Lütkepohl (2006, p. 611), this two-layer property can be expressed as the dependency of an observable and possibly multiple time series $y_1, ..., y_\mathrm{T}$ upon an unobservable state β_t that is driven by a stochastic process, whereas the dependency between y_t and β_t is described in a measurement equation that takes the form of the TVP VAR in our specification (explained below).

In this paper, time-variation is introduced in a way that allows the policy propagation parameters – but not the structure of the error variances covariances – to potentially evolve over time as observations are added, i.e. the analyses are based on a homoscedastic TVP VAR that generates the measurements.[28] Some general properties of the applied n-dimensional TVP VAR of order p including an intercept term are presented below. For each observation t a standard time invariant VAR model introduced in part 4 can be rewritten in time-dependent, stacked and vectorized form,

$$\mathrm{y_t} = Z_t\beta_t + \varepsilon_t, \quad t = 1, ..., \mathrm{T}, \tag{1}$$

where the data vector y_t and the residuals ε_t are of dimension $n \times 1$, and the (independent) residuals follow a zero mean process with **time-invariant residual variance covariance** matrix $E\left(\varepsilon_t\varepsilon_t^{'}\right) = H_t = H$. The $m \times 1$- vector β_t contains the values for the n constants and pn^2 lag coefficients in period t. It is derived from vectorization of the $n \times 1 + np$ dimensional stacked form coefficient matrix, $\beta_t = vec\,(B_t)$, where $B_t = [\nu_t : A_{1,t} : ... : A_{p,t}]$ includes the $n \times 1$ vector of constant terms ν_t and the $n \times n$ matrices A_i, for $i = 1, ..., p$ of the lag coefficients. The regressor matrix is not restricted and only contains a time-varying constant as an exogenous explanatory variable. It is defined as $Z_\mathrm{t} = \left(Z_{t-1}^{'} \otimes I_n\right)$ and is of dimension $n \times m$, where $Z_{t-1} = \begin{bmatrix} 1 & y_{t-1} & ... & y_{t-p} \end{bmatrix}'$ is the $1 + pn \times 1$ regressor matrix of the stacked

[27] A wide range of alternative specifications have been suggested, including Markov-switching VARs (e.g. Paap and van Dijk (2003), or Sims and Zha (2006)) and other regime-switching VARs (e.g. Koop and Potter (2006)).

[28] We will not consider time-varying moments of the processes, covered, for example, in cointegration approaches based on trended series. Instead, we will continue working with the de-trended and seasonally adjusted series already used in the first sections.

form TVP VAR. A sample size of T time series observations as well as p presample values for each variable are available. With regard to the dynamic path of the $m \times 1$ state vector β_t we will follow the economic literature on state-space models where the states are generally assumed to be explained by an (m-dimensional) first-order autoregressive process (see, for example, Hamilton (1994), Doan, Litterman and Sims (1984) or Harvey (1992)) as represented in the following state (or transition) equation (2),

$$\beta_t = T_t \beta_{t-1} + c_t + R_t \eta_t, \quad t = 1, ..., \mathrm{T}, \tag{2}$$

where the $m \times m$ matrix T_t is used as the transition matrix that reflects all the information on how past states enter the measurement equation at time t, the $m \times 1$ vector c_t consists of other exogenous components such as a constant or dummy variables, R_t is a matrix of dimension $m \times g$ that involves structural relations between the disturbances of the states that are described in the $g \times 1$ vector η_t and that have a **time-invariant** $m \times m$ **variance-covariance** matrix, Q. Moreover, Q, T_t, c_t and R_t are referred to as the state system matrices.
The following four assumptions are made to keep the structure of our state-space model as simple as possible

1. We already implemented linearized relations in the measurement and state equation.

2. We further assume the measurement errors and the disturbances of the state equation each to be serially uncorrelated, i.e. $E(\varepsilon_t \varepsilon_s') = 0$, $E(\eta_t \eta_s') = 0$, for all $s \neq t$, and - as we have mentioned - to follow a multivariate normal distribution with mean zero and time-invariant variance-covariance matrices H and Q. Furthermore, they are assumed to be uncorrelated with each other and with some initially chosen normally distributed state variables, β_0, i.e. $E(\varepsilon_t \eta_s') = 0$ for all t and s, and $E(\varepsilon_t \beta_0') = 0$, and $E(\eta_t \beta_0') = 0$ for $t = 1, ..., \mathrm{T}$. The last two assumptions also guarantee no correlation between ε_t and β_t as well as no correlation between η_t and β_{t-1}, $E(\varepsilon_t \beta_t') = 0$ and $E(\eta_t \beta_{t-1}') = 0$ for all t.[29]

3. Besides the time-invariant system matrices H and Q, we further assume the other system matrices to be time-invariant, $Z_t = Z$, $T_t = T$, $c_t = c$ and $R_t = R$, where $R = I_m$ as g equals m.

4. In some of the most recent TVP VAR studies, VAR coefficients are set to evolve according to random walks (see Koop and Korobilis (2009)). For a transition factor equal to an indentity matrix of dimension m, $T = I_m$, the state equation in (2) is transformed into a random walk. To us, this seems reasonable based on the properties of the economic relations between the fiscal stance and the output gap. Due to the unit root (or I(1))-characteristics of such a process, the states will follow a stochastic trend, either a trend in a positive or a negative direction. With regard to fiscal policy issues, this is exactly what we expected in our hypothesis of gradually changing regimes. Another advantage of such random walk models is that they are typically well-suited for forecasting macroeconomic

[29] The normality assumption of the disturbances is required because we will use the standard linear recursive algorithm of the Kalman Filter, which is derived on the basis of density functions to estimate the states (see Tanizaki (1996)).

time series.[30]

Based on these specifications, we will estimate the normal linear state-space model in the next step.

5.2 Maximum likelihood estimation of the state space model - A five step procedure

Our estimation strategy is based on conventional methods for models of econometric time series. At its core is the evaluation of the log-likelihood function using Kalman filtering recursions over a set of hyperparameters that contain randomly drawn or (partly) predefined variances-covariances and some *a priori* defined values in order to find the states that maximize the log-likelihood. Basically, the procedure is based on the following five step algorithm.

1. Randomly drawing of some initial variances-covariances for the state and measurement residuals.

2. Application of the Kalman filtering recursions using the information from the initially drawn variances-covariances and some defined state priors to estimate the states and their variances-covariances.[31]

3. Evaluation of the log-likelihood of the estimated state space model.

4. Drawing of new variances-covariances for the state and measurement residuals close to the previous ones, estimation of the states and computation of the corresponding log-likelihood.

5. Repetition of the whole procedure until the log-likelihood maximized for the corresponding estimates gives us the optimal posterior states and their posterior variances-covariances.

The evaluation of the log-likelihood function is obviously of special importance. It is evaluated conditional on the available data $y = (y_1, ..., y_\mathrm{T})'$ applying an iterative algorithm to find the optimal estimators for the residual variances-covariances of the measurement equation, H^*, and - depending on the scenario - also of the state equation, Q^*, given some initial residual variance-covariances, H and Q as well as some prior settings $\beta_0 \sim N(\mathrm{b}_0, \mathrm{P}_0)$. A derivation of the log-likelihood function for our specific estimation problem that is based on Kalman filtering recursions can be found in appendix C.
Kalman filter recursions are useful tools within the evaluation process of the likelihood function because they provide – given the measurements $y_1, ..., y_\mathrm{T}$ and under normality assumptions – an optimal estimator b_t for the states β_t at each period of time. The filtering estimates that result from the updating step at each point of time t are given by the expected value of the state variable, conditional on the observations up to this point of time, $\mathrm{b}_t = \beta_{t|t} = E(\beta_t | y_1, ..., y_t)$ for $t = 1, ..., \mathrm{T}$. Once our iterative algorithm converges to an optimum, H^* and Q^*, the corresponding Kalman

[30] In addition, we experimented with other versions that are common in macroeconometrics. We used mean reverting versions of equation (2) that can be derived by setting $0 < |T| < I_m$ and the constant term equal to the weighted steady state or mean of the state vector. For choices of $T = 0.99$ or $T = 0.999$, which corresponds to the values chosen by Muscatelli et al. (2007), no significant differences can be found.

[31] At this iterative step a stationarity test and (if needed) a correction mechanism are implemented.

filter recursions represent the optimal estimates for the posterior states, b_t^* and their posterior variances-covariances, P_t^*.[32] However, to start the Kalman filter recursion algorithm some initial values for the states and their variances-covariances have to be specified *a priori*.

5.3 Initialization: Prior state distribution and sensitivity analyses

To start the Kalman filter recursions, we defined a prior distribution for the states described by a prior mean and a prior variance-covariance matrix. Moreover, in the less comprehensive scenarios we will concentrate on (in which the overall number of hyperparameters at each step of the iteration is systematically reduced by exogenous choice of the residual variance covariances of the states) the m diagonal elements of the residual variance covariance matrix Q have to be specified *a priori* as well.[33] As there are different options and the initialization of the underlying system has an effect on the whole estimation procedure and the posterior states, we studied the impact of different approaches in great detail in sensitivity analyses. In doing so, the exogenous choice of the states residual variance covariance matrix turned out to provide the biggest leverage on our results. Therefore, it has to be chosen with particular care.

5.3.1 Specification of the state-residual variances covariances

Some basic relations can be observed in a sensitivity analysis. The smaller the elements on the main diagonal are set, the less fluctuation in the residuals of the states is allowed. And thus, the less (gradual) evolvement can be covered by the reduced-form coefficients of the AR(1) state equation, because, in such a case, the state coefficients of any period can differ from their lagged values only by small amounts. In other words, the time-varying system approximates a system with constant parameters more closely. Moreover, there is an important interaction at work: Any potential evolvement that could have been covered by more varying parameters is now directly transferred to the residuals of the measurement equation. The estimated system moves further away from actual data.[34]

Obviously, there exists a serious *trade-off between fitting the data and the implementation of an economic structure.* Setting the variances too low would lead to the implementation of higher degrees of economic structure but diminish any potential explanatory gain resulting from time-varying parameters. In contrast, setting the diagonal elements too large would lead to a significant loss of economic structure explained by the estimated system because the system then just reproduces the evolvement of the actual data.

To be able to fine-tune the "looseness of the states" or the "tightness of the state residuals", we implement a multiplicative leverage on an identity matrix. And we

[32] We used the slightly modified method of Anderson and Moore (1979) to compute the variance-covariances of the filtered states in order to avoid negative definite matrices because of round-off errors in the Kalman filtering recursions.

[33] In this paper, we focus on versions where only the variance covariance matrix of the measurement errors depends on some randomly drawn and normally distributed parameters and has to be estimated. In fact, exogenous identification seems to be reasonable because it helps to mitigate serious problems associated with the proliferation of parameters to be estimated in a TVP VAR approach. According to Koop and Korobilis (2009), it is hard to obtain precise estimates of coefficients in models where the number of coefficients to estimate is T times higher than in constant-coefficient models. Impulse responses can have dispersed posterior distributions leading to wide confidence intervals. Moreover, the system is in danger to be overparameterized.

[34] Systematically larger disturbances decrease the degree of explanation of the estimated measurement equation.

decided to put more weight on the implemented degree of economic structure than on fitting the data, which seems to be reasonable given that the main focus of this study is on structural relations and not on forecasting. Therefore, the leverage λ is finally set rather tightly and equal to 10^{-9} in the benchmark scenario after testing a wide range of reasonable values.[35]

5.3.2 Prior distribution of the states: Training sample priors

The choice of the prior distribution of the states also leaves some degrees of freedom to the analyst. In line with existing studies, we use some LS estimations to set the prior distribution of the states as a benchmark (see, for instance, Primiceri (2005)) or Koop and Korobilis (2009)). In doing so, the prior state vector is calibrated based on the least squares point estimates of a time-invariant VAR(p) process with the same lag order p as for the time-varying parameter VAR and for a training sample of 20 quarters starting in the first quarter of 1970. This leads to $\beta_{0|0} = \mathrm{b}_0^{\mathrm{LS_{TS}}}$ as the prior mean. Concerning the conditional variance-covariance matrix $P_{0|0}$, simply using the training sample LS estimates may be too strict and thus reduce the reliability of the estimated model. To keep it simple, we calibrated the $m \times m$ prior variance-covariance matrix on a m-dimensional diagonal matrix with the corresponding training sample LS estimates on the main diagonal and multiplied this matrix by another regulating leverage called τ ("tightness of the state coefficients"), which postulates our view on how binding the prior expectations for constant terms and lag coefficients in the state vector are for the posterior states.[36] To reduce the impact of the prior mean or the degree of prior information that is used to compute the posterior states, we set τ equal to 4 which is consistent with Primiceri (2005), Baumeister, Durinck and Peersman (2008) and Koop and Korobilis (2009) in their Bayesian TVP-VAR approaches. Summing up, the initial state vector is assumed to be normally distributed with the conditional expected value $\mathrm{b}_0^{\mathrm{LS_{TS}}}$ and the conditional variance-covariance matrix $P_{0|0} = \mathrm{P}_0 = \tau \cdot \mathrm{P}_0^{\mathrm{LS_{TS}}}$,

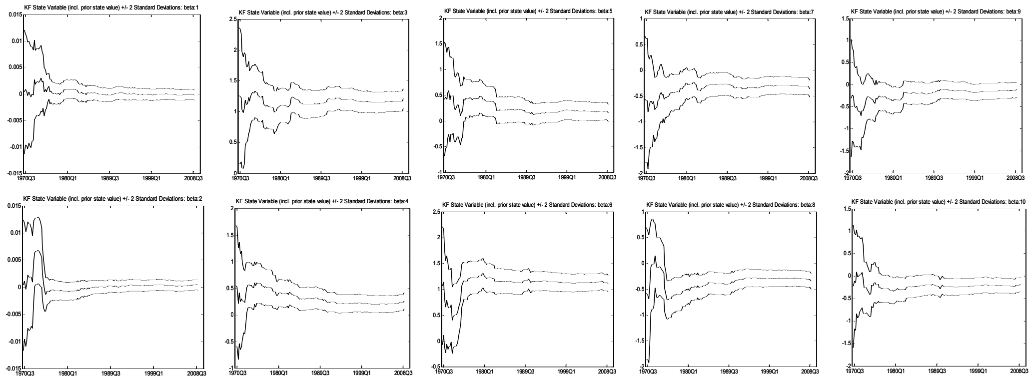
$$\beta_0 \sim N\left(\mathrm{b}_0^{\mathrm{LS_{TS}}}, \tau \cdot \mathrm{P}_0^{\mathrm{LS_{TS}}}\right). \tag{3}$$

To make sure that our posterior results do not depend too much on the setting of the prior distribution, we additionally apply another much more diffuse prior distribution that reflects little or no information regarding the values of the unknown parameters. In two different versions we implemented a vector of zeros and full sample least squares estimates as the prior mean in combination with a very diffuse prior variance-covariance matrix that is now independent of the sample size, $P_{0|0} = \mathrm{P}_0 = \tau \cdot I_m$ with a tightness factor on the state coefficients equal to 10^{-5}. The estimation results

[35]We check for the sensitivity of the results for different choices of λ within a reasonably bounded space. An upper bound is derived from a comprehensive estimation where both, the residual variance covariance matrix of the measurement equation and the states residual variances are estimated by ML. For such large estimations the probability of overparameterization increases dramatically and the estimations can be heavily biased by outliers that are included. We expect rather high values, that serve as some kind of upper bound. This value is around 10^{-3}. A lower bound is derived based on the existing literature. Applying the Minnesota priors the way that Muscatelli et al. (2007) implemented them – the authors used 10^{-7} times the LS-estimated variance of the states – corresponds to a looseness parameter around 10^{-9} in approaches where Q does not depend on any LS-estimated state variances.

[36]For higher values of τ the initially chosen states become less binding for the posterior states and, thus, the authors confidence in the priors is lower.

Figure 4: Optimal posterior states



of both non-informative initializations are very similar to the estimates based on the training sample and they do not lead to totally different or even switching impulse responses, compared with the results from the reference scenario. Thus, the time-varying structure seems to be robust with respect to different prior specifications.[37]

5.4 Estimation results: Posterior states

In the reference approach, only the variances-covariances of the measurement equation have to estimated because the state-covariance matrix is determined *a priori*. The optimal estimates for the time-varying state variables are presented in figure 4 within two-standard deviation bounds.

Comparing the estimated time-invariant measurement residual variances and covariances of the benchmark TVP models with the estimates from time-invariant and time-invariant regime-switching vector autoregressions (all presented in table 1 in appendix D), the differences are only very small.[38] However, compared with TVP scenarios in which we allowed for stronger fluctuations in the states (Appendix D, table 2), these variances-covariances are rather large – which we expected because we decided to put a higher weight on economic structure than on fitting the data. Nonetheless, regarding the coefficients of determination, the goodness of fit of the TVP VAR under our "conservative" benchmark specification is larger for each single equation, which is confirmed by the Schwarz information criterion (SIC), whereas, based on the Akaike information criterion (AIC), the two constant parameter models perform better. Nevertheless, this inconsistency of the criteria can be eliminated by simply putting only slightly more weight on fitting the data in the choice of the priors.

5.5 Structural analyses

To analyse the dynamic fiscal impacts of shocks in the business cycle we rely again on forward-looking fiscal policy reaction functions. Now, these systematic reaction

[37]Additionally, we used Minnesota priors as they offer some more useful leverages to manage the prior specification. However, for common specifications of the leverages (see Muscatelli et al. (2007)) we generated results similar to the ones based on presample LS priors and non-informative LS-based priors. Estimation results and the corresponding impulse responses are available from the authors upon request.

[38]Standardized measurement residuals are presented in appendix D.

functions can evolve gradually over time in their sensitivity and pattern, except for the contemporaneous period, because the estimated measurement residual covariance matrix is homoscedastic.[39] We implement the same time-invariant B-SVAR identification scheme as introduced in section IV for the case of time-varying parameters. Thus,

$$\mathrm{B}\varepsilon_t = u_t, \tag{4}$$

where $G = E\left(u_t u_t'\right) = \mathrm{B}E\left(\varepsilon_t\varepsilon_t'\right)\mathrm{B}' = \mathrm{B}H^*\mathrm{B}'$ and H^* is the optimally estimated ML variance covariance matrix of the measurement residuals. B^{-1} is a $n \times n$ lower triangular Cholesky-type decomposition matrix with unit diagonal, such that G is an n-dimensional diagonal matrix with the variances of isolated or respectively orthogonalized measurement errors that now indicate economic innovations on the main diagonal.

Because point estimates or any transformed version of them might not be reliable, we additionally applied bootstrapping procedures (2,000 replications) to simulate artificial distributions that are used to compute bootstrapped standard deviations and confidence intervals for the impulse responses (figures shown in appendix E). We use (data-based) non-parametric techniques as serious problems can arise in time-varying parameter approaches, applying standard parametric bootstraps that do not adjust for the scale and distribution of the estimated measurement residuals.[40]

5.5.1 Results

On the basis of these specifications, we can now derive the effects of output gap shocks on the primary balance if gradual evolvement of fiscal policy regimes is possible.

We start with a discussion of the time-varying impulse response functions. These show the impact of an output gap innovation in figures 5a-5f. As in the case of the time-invariant benchmark model, the development of the output gap itself (reflecting the state of the business cycle) has to be taken into account when the effect of the output innovation on the primary balance is traced through the endogenous system. Figures 5a-5c show the effect of a 1 percentage point (PP) output gap shock on the output gap itself and Figures 5d-5f the reaction of the primary balance to the same shock (both reactions in PP of GDP), starting in 1976 after a burning-in period of 24 quarters.[41]

In a first step, we discuss **the effects of output gap shocks on economic development** (Figures 5a-5c, which all show the same reactions from different perspectives).

[39] In cases of heteroscedastic residuals, different contemporaneous reactions can be generated by a Cholesky-type decomposition - if (for example) the volatility of the measurement residuals is assumed to follow a certain stochastic process (see, for example, Primiceri (2005) or Baumeister, Durinck and Peersman (2008)).

[40] We have mentionend that the scale of the estimated measurement residuals heavily depends on the prior choice of the "tightness of the state residuals". In the standard parametric procedures the bootstrap residuals used to simulate the actual measurement errors are always pseudo-random values drawn from a standard normal distribution and thus, 95% of the standardized bootstrap residuals are, by definition, located within values of approximately +/- 2. Hence, the parametric bootstrapping procedure would generate reliable results only in cases of rather tight leverages, where the actual measurement residuals are as high as the bootstrap residuals.

[41] As an alternative to the point estimates, we have estimated the median impulse response functions (not reported here) – which show similar developments from strong reactions to less strong reactions. We decided to focus on the optimal point estimates as the estimation of median responses in the time-varying model leads to a strong smoothing and makes the specific periods of regime changes harder to detect.

The reaction in the contemporaneous quarter of a shock is necessarily always one. Within the first three quarters after a shock, the reaction of the output gap to an output gap shock differs only very slightly over the period analyzed (1976-2008). After the contemporaneous shock of 1 PP, output increases further in the first quarter after the shock. This reaction is very stable and fluctuates around 1.20 percentage points throughout the sample. The fluctuations around that level are somewhat larger before 1991 than thereafter (especially in the 1970s), but altogether they are small. In the second quarter, the reaction stays on a level which is similar to that in the first, before it falls back to the initial level of around 1 PP in the third quarter.

From quarter four after a shock onwards, the reaction to the shock continues to fade out, but the variation over time increases. Here, we can distinguish four periods.

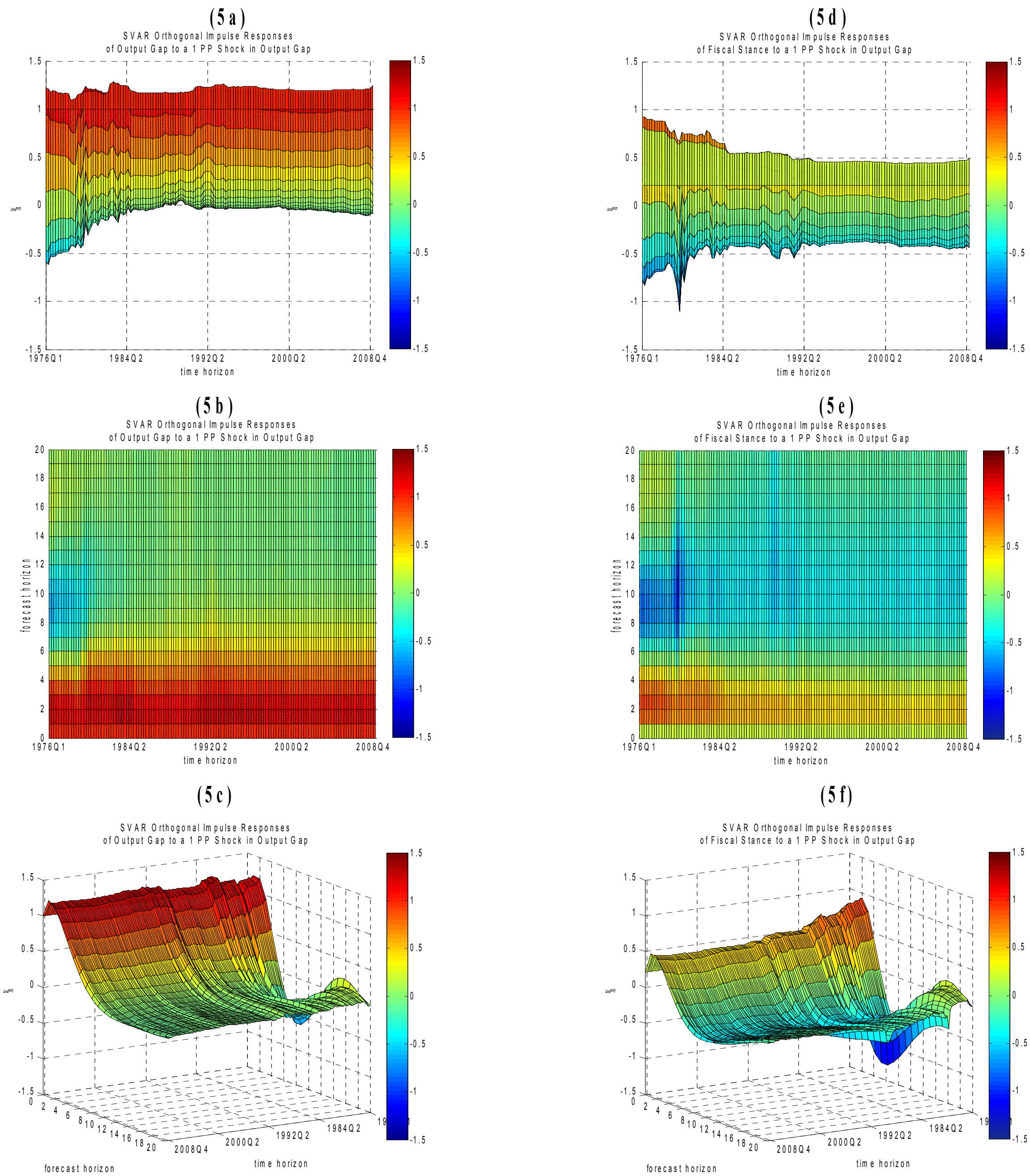
In the first period from 1976 to 1980, the output gap shock fades out very fast. In the fourth quarter after a shock, the gap is already reduced to 0.5 PP and within the sixth quarter it has already turned negative with -0.2 PP. In the following quarters it even falls to -0.5 PP before the output gaps adjusts to its zero equilibrium level again. In the second period from 1980 to 1984, the fading out of the positive output gaps slows down gradually and the zero line is not passed before quarter eight. Additionally, the output gap does not turn substantially negative in later quarters any more. From 1984 onwards, the pattern becomes very stable with gaps that slowly fade out and are positive even until quarter ten. The period from 1984 to 1991 is characterized by some gradual volatility – especially towards the end of the period – but the general pattern looks similar to the one we observe after 1991. In the last period from 1991 to 2008, we observe a reaction of output to output gap shocks, which is very similar to the one in the time-invariant model and is characterized by the build-up of an output gap in the first two periods after a shock (to around 1.2%) and a gradual reduction down to zero by the tenth quarter (2.5 years after a shock). The most important difference to the results of the time-invariant model (see part 4) seems to be that the output gap turns only very slightly negative and does so only towards the end of the period analyzed (in the late 2000s).

We can therefore summarize the following: The reaction of output gap in the first three quarters after an output gap shock is very similar throughout the time period analyzed, while we can distinguish four different regimes with respect to the reactions after the third lag. These regimes differ strongly with respect to the speed of the fade-out of the output gap shock and with respect to how negative the output gap reaction becomes around 2 years after a shock.

Against the background of the cyclical development of GDP, we can now discuss the **reaction of the primary balance to the output gap shocks** (see figures 5e-5f for the time-varying responses to a 1 PP output gap shock). We see a contemporaneous reaction of the primary balance to changes in the GDP gap, which is likely to reflect the working of automatic stabilizers as discretionary fiscal policy is hardly able to react to economic shocks within the same quarter (see discussion in part 4). The effect equals 0.21 percentage points – nearly the same value as in the time-invariant model (0.22 PP). With respect to the reactions thereafter, we can distinguish different reaction patterns in the four same periods (or regimes) we discussed with respect to the output gap (1976-1980, 1980-1984, 1984-1991 and 1991-2008).

At the beginning of the **first regime** from 1976 to 1980, we see that the primary balance continues to increase in the first quarter (reaching 0.81 PP) and even further in the second quarter after a shock (reaching around 0.92 PP). This development is accompanied by a positive output gap reaction (see figures 5 c-e), which increases further in the first quarter after the shock and then stays on a similar level in the

Figure 5: Time-varying impulse responses



second quarter after a shock. It thus reflects a countercyclical policy reaction. In the following quarters, the reaction of the primary balance decreases – nearly parallel to the decrease of the output gap. Five quarters after the shock, the fiscal stance turns negative – at the same time as the output gap – still indicating a countercyclical fiscal policy reaction. During the next quarters, the fiscal stance turns strongly negative (to a level of around -0.79 PP) while the output gap decreases to, on average, -0.50 PP. From the start point in 1976 Q1, these countercyclical reactions are gradually reduced – with respect to the positive reaction in the first quarters as well as with respect to the negative reactions later on. By 1979 Q2, the maximum positive reaction has been reduced to 0.79 PP and the fiscal stance is then reduced later after a shock to only -0.60 PP.

The **second regime**, covering the second oil crisis and lasting until 1984, is a gradually evolving fiscal policy regime as well. Its general pattern looks similar to the first regime. Here, the fiscal stance initially increases slightly less strongly than in the first regime and reaches only 0.64 PP (one quarter after a shock) and 0.72 PP (two quarters after a shock) in 1981 Q4. It should be noted that this reduction in the countercyclical reaction in the two quarters after the shock cannot be explained by differences in the output dynamics, as the output gap reacts (when compared to the first regime) more strongly in the second regime (we will come back to this point later). The fiscal stance is reduced to zero in the fifth quarter after a shock along with the output gap, which also reaches zero five quarters after a shock. In contrast to the first regime, the downturn in the output gap is less severe and the fiscal stance attains only negative values of on average around -0.45 PP close to the tenth quarter after an output gap shock. According to our definition, this again reflects a countercyclical reaction of the fiscal stance throughout the years after a shock as positive (negative) fiscal balances are accompanied by positive (negative) output gaps. At the end of this regime, just before transition into the next regime starts (1983 Q4), the positive reactions in quarters one and two after a shocks are reduced to 0.60 PP and 0.66 PP respectively.

In the **third regime (1984-1991)**, we see nearly no evolvement of the responses. But compared to the second regime we see a further reduction of the fiscal stance reaction. In 1984 Q4 the reaction in the first quarter after the shock equals around 0.56 percentage points in this regime until the transition to pan-German data in 1991 induces some moments. This might be explained by the slightly lower reaction in the GDP gap in the first quarter after a shock in this regime. However, another important observation is that the fiscal stance no longer increases in the second quarter after the shock but slightly decreases – which does not seem to be directly related to the reaction in the output gap. In the third regime, the fiscal stance is roughly -0.45 PP around the tenth quarter. This takes place, although the output gap shock nearly fades out without turning negative. The result is a fiscal policy reaction, which can still be called countercyclical. However, as the output gaps are very small from around the seventh quarter after a shock, it can be classified alternatively as acyclical fiscal policy from the seventh quarter onwards.

In the **fourth regime (1991-2008)**, we observe a further reduction in the countercyclical reaction of the fiscal stance in quarter one and two after the shock (to 0.48 PP and 0.46 PP respectively) – despite the fact that the reaction in output to output shocks in these quarters is larger than in the previous regime. The reductions of the output gap in the following quarters are accompanied by a reduction in the fiscal stance and, as both are positive till around the seventh quarter, this again reflects countercyclical policy. From the seventh quarter after the shock, we see deficits in this regime, while the output gap stays very slightly positive. Generally, this would mean that we observe procyclical fiscal policy but, as the output gaps are very small, these periods might again be classified as acyclical fiscal policy. At the end of the sample in 2008 Q4, the maximum of the fiscal stance response has been reduced to 0.46 PP of GDP reached in the first quarter (compared to 0.81 PP in the first regime). Deficits reach -0.40 PP of GDP around quarter ten.

How can we summarize our findings? Throughout the sample (1976-2008), we observe strong countercyclical fiscal policy reactions in the first quarters after an output gap shock. This reaction decreases over the different regimes.[42] As the reaction of output to the output gap shock in the first quarters after a shock does not vary much over our data set, this change is likely to result from changes in fiscal policy (be it in the form of the automatic stabilizers or of discretionary fiscal policy) over time. In the first and the second regime, we observe a countercyclical fiscal policy reaction throughout the 20 quarters (5 years) after a shock analyzed here. The output gap turns strongly negative around the tenth quarter after a shock and the fiscal stance reacts with primary deficits at the same time. In the third and fourth regimes, the output gap does not turn strongly negative in later quarters after a shock but the primary balance continues to do so. This makes the fiscal policy reaction in later quarters, at best, acyclical.

6 Conclusion and outlook

The timing of fiscal policy has an important impact on the effectiveness of fiscal stimulus programmes but has received only limited attention. Studies that take the endogeneity between economic shocks and fiscal reactions serious are rare, and most analyses choose a time-invariant perspective. In this paper, we have pursued a time-varying VAR analysis of the interplay of economic shocks and the reaction of public finances in Germany from 1970-2008.

In our analysis, we focus on overall fiscal policy – not at last because an reliable time-varying method of cyclical adjustment for quarterly data is still to be established. To be able to discuss the interplay between economic shocks and the fiscal stance adequately, we analyze the economic effects of output shocks in parallel with their effects on public finances.

In the time-invariant model for quarterly data from 1970-2008, we find a countercyclical reaction of the fiscal stance to output gap shocks in the first four quarters after a shock and an acyclical reaction afterwards. However, parameter stability tests point strongly to instability and the inadequacy of time-invariant models.

We therefore applied a time-varying parameter (TVP) VAR to the same data. The suspected time-varying regimes with regard to the reaction of output gaps to output gap shocks and of the timing of fiscal policy over the business cycle were strongly confirmed by this approach, as we found four distinctive regimes: **1976-1980**, **1980-1984**, **1984-1991** and **1991-2008**. The first implication of this finding is that changes in the fiscal policy reaction became rarer over time as the time span covered by the regimes tended to increase.

In all regimes we continue to observe a strong countercyclical reaction of the fiscal stance in the first quarters after an output shock. Nevertheless, this reaction is reduced steadily from regime to regime and falls from around 0.9 percentage points

[42] This finding is consistent with other studies such as and Aghion and Marinescu (2008), who find a decreasing reaction in the first year after a shock for euro-area data from 1990 to 2005. Further support for German data comes from Debrun and Kapoor (2010).

of GDP in the first regime (maximum reaction two quarters after a shock) to slightly less than 0.5 PP (maximum reaction one quarter after the shock) in the fourth regime. This finding is generally consistent with the results of Aghion and Marinescu (2008) for the development of the cyclical response of the fiscal balance to output shocks in EMU countries from 1990 to 2005. The relative stability of the output gap reaction to output gap shocks shows that this development of the fiscal policy reaction does not stem from changes in the economic reactions (which could be interpreted as a change in the economic volatility) but rather that the reasons lie in the field of public finances (either with respect to automatic stabilizers or with respect to discretionary fiscal policy). In the second and third years after a shock, the timing of fiscal policy differs as well over the regimes. In the first regime (1976 to 1980) and, to a slightly lesser extent, in the second (1980-84) regime, the output gap turns strongly negative in the second and third years after an output gap shock. In these regimes, positive primary balances are accompanied by positive output gaps in the first two years and negative primary balances are accompanied by negative output gaps in later years after a shock, which indicates countercyclical fiscal policy in all years after the shock. During the third (1984-1991) regime and the fourth (1991-2008) regime, however, the output gap turns only very slightly negative in the second and third years after the shock. However, the primary balance continues to turn negative in the second and third years after a shock. Hence, the pattern changes in later years after a shock to an acyclical timing of fiscal policy. For the third and the fourth regimes, the output gap reaction had changed, while the fiscal policy reaction changed far less – which might indicate that, again, some structural change in public finances – be it in the form of automatic stabilization or of discretionary fiscal policy – has taken place.

Based on these findings, we can offer some important insights: overall fiscal policy in Germany from 1970 to 2008 was strongly, but, over time, decreasingly countercyclical in the first four quarters after a shock. In the first two regimes (till 1984), fiscal policy was even consequently countercyclical in the five years after economic shocks analyzed here, while the reactions of the fiscal stance were, first, countercyclical and, then (from around 2 years after a shock), acyclical in the third (1984-1991) and fourth (1991-2008) regimes. The comparison of output and fiscal reactions point to major structural changes in public finances between the regimes – with respect to the reaction in the first year as well as in the second and the third years after a shock.

Interesting questions, which follow from these findings, are: "Which structural changes are responsible for the different fiscal policy reactions under the different regimes?"; "Have the automatic stabilizers changed, for example, owing to changes in the progressivity of the tax system or, say, because of reforms in the unemployment benefit system?"; "Or was it the timing of discretionary fiscal policy which differed under the four regimes?" We plan to tackle these questions in future analyses, which apply the methodology of this paper to data, which are adequately cyclically adjusted.

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Appendices

Appendix A: Excursus - A more detailed look at the automatic stabilization channels at work

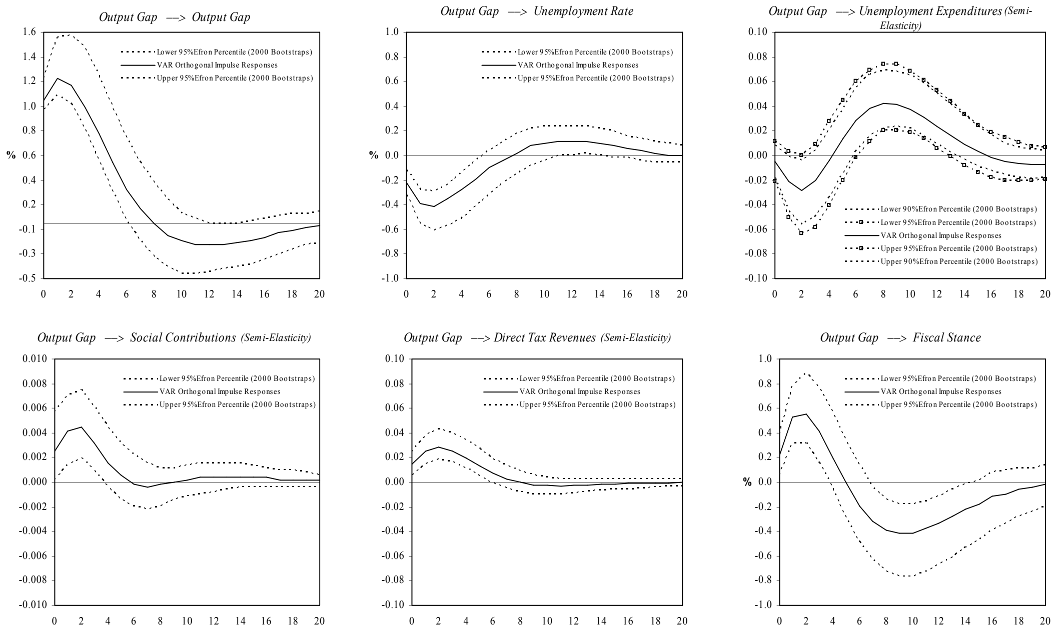


Figure a: Impulses responses of cyclical components of nemployment rate, log unemployment expenditures, log social contributions, log direct tax revenues and fiscal stance to a 1 percentage point output gap shock

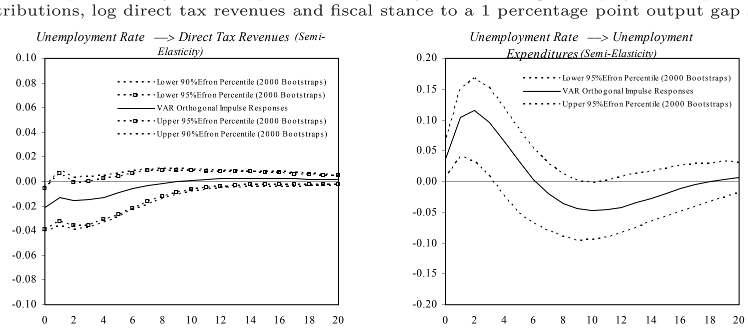


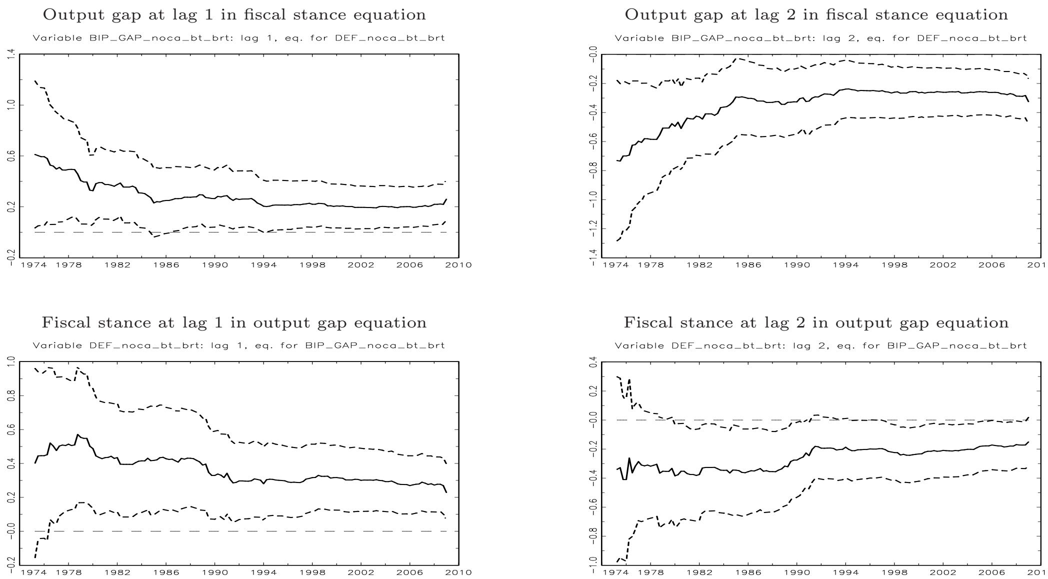
Figure b: Cross-Checks: Impulse responses of direct taxes and log unemployment expenditures to a 1 percentage point shock in the unemployment rate

Studies of automatic stabilization (see, for example, Girouard and André (2005) or Mohr et al. (2001)) identify direct tax revenues and unemployment expenditures as the main channels through which automatic stabilizers work. We add an analysis based on higher dimensional VARs to control for the importance of automatic stabilization in our set-up.[1] Our estimations show that the direct tax channel seems to dominate the creation of the initial primary surpluses after an output gap shock,

[1]The corresponding impulse responses of the unemployment rate (as a percentage = percent deviation/100), the log unemployment expenditures and the log direct tax revenues to a 1 PP output gap shock are based on a six-dimensional VAR(2), which furthermore includes the component of log social contributions and the fiscal stance. Identification of the output shock is achieved by implementing a recursive Cholesky-type structure of the economy. The results are not significantly sensitive to the ordering of the variables. The same recursive scheme in combination with the same ordering of the variables is used to identify 1 PP shocks in the unemployment rate.

whereas the unemployment expenditure channel dominates the creation of the deficits following the period of declining output gaps in the overall fiscal stance adjustment process to an output gap shock.

Appendix B: Recursive least squares estimates of VAR with additional deterministic



Appendix C: The log-likelihood function

The log-likelihood function of the underlying Gaussian model can be derived in three steps.[2] In a first step, using Bayes' theorem and the sample density function, the joint density function can be derived, where $\theta = (H, Q, \mathrm{b}_0, \mathrm{P}_0)$ is the vector of hyperparameters and $\mathrm{f} = f\left(y_t \left| Y_{t-1}\right.; \theta\right)$ the distribution of y_t, conditional on the information set at time $t-1$, Y_{t-1}. Additionally taking the Gaussian properties of our model into account, the true state vector of our state space model at time t is normally distributed by assumption with mean $\mathrm{b_t}$ and variance-covariance matrix $\mathrm{P_t}$. Therefore, also y_t is normally distributed with recursively given mean $E\left(y_t \left| Y_{t-1} \right|\right) = y_{t|t-1}$ and variance-covariance matrix $Cov\left(y_t \left| Y_{t-1}\right.\right) = F_t$. In a second step, the Kalman filtering equations can be used to estimate these quantities, given a specific vector of hyperparameters $\theta = (H, Q, \mathrm{b}_0, \mathrm{P}_0)$. Thus, $y_{t|t-1} = Z\mathrm{b}_{t|t-1}$ and $F_t = Z\mathrm{P}_{t|t-1}Z' + H$. Finally, taking the joint density function, the normality assumption and the information from the Kalman filtering recursions, the log-likelihood function of our Gaussian state space model is,

$$\ln L\left(\theta \left| y\right.\right) = \ln L\left(H, Q, \mathrm{b}_0, \mathrm{P}_0 \left| y\right.\right) = -\frac{\mathrm{nT}}{2}\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{\mathrm{T}} \ln|F| - \frac{1}{2}\sum_{t=1}^{\mathrm{T}} \nu_t' F_t^{-1}\nu_t,$$

where we denote the dimension of our data matrix by a general dimension n, and ν_t is the estimation error from the Kalman filtering procedure. This expression is maximized with respect to the vector of hyperparameters $\theta = (H, Q, \mathrm{b}_0, \mathrm{P}_0)$, where the factors $\mathrm{b}_0$and P_0 are the initially set priors and thus, only H, and – depending on the scenario – Q has to be estimated as well. Obviously, this problem is an

[2]Our derivation is based on Lütkepohl (2006).

Table 1: Estimates of the Measurement Residual Variances Covariances and Model Statistics (Sample: 1970 Q3 – 2008 Q4)

	a) TVP VAR	b) Regime-Switching VAR	c) Benchmark VAR
Parameter and Model Statistics	ML Estimates, $\lambda = 10^{-9}$, Training Sample Priors)	LS Estimates	LS Estimates
$\hat{h}_{11}^2$	$0.2749 * 10^{-4}$	$0.2400 * 10^{-4}$	$0.2752 * 10^{-4}$
$\hat{h}_{22}^2$	$0.2351 * 10^{-4}$	$0.2351 * 10^{-4}$	$0.2354 * 10^{-4}$
$\hat{h}_{12}^2$	$0.0586 * 10^{-4}$	$0.0584 * 10^{-4}$	$0.0582 * 10^{-4}$
$R^2_{adjusted}$(Eq. 1)	0.8611	0.8603	0.8354
$R^2_{adjusted}$(Eq. 2)	0.9070	0.8965	0.8935
Log Likelihood	1175.20	1216.77	1201.30
Determinant (Cov)	$0.6121 * 10^{-9}$	$0.4702 * 10^{-9}$	$0.5748 * 10^{-9}$
AIC	-21.111	-21.244	-21.147
SIC	-20.955	-20.889	-20.950

unconstrained nonlinear optimization problem that can be solved using a standard iterative algorithm.[3] In fact, the procedure we implemented is based on the dual approach of the optimization problem, the minimization of the negative log-likelihood function.

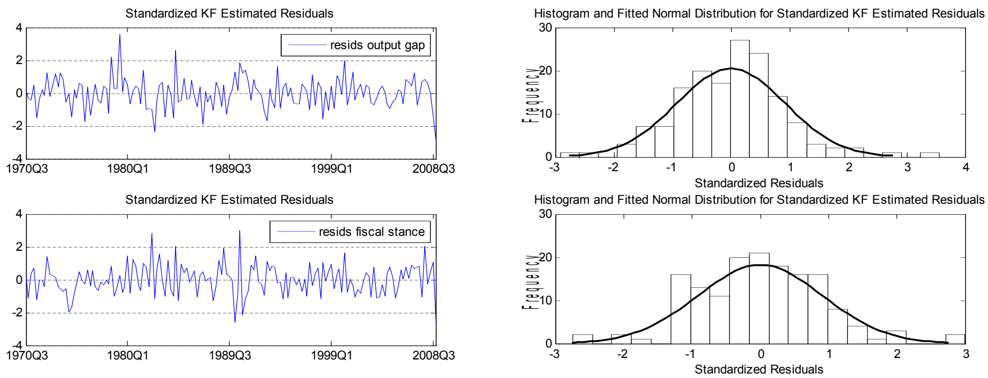
Appendix D: Estimation Results

Table 1 shows the ML estimates of the parameters of interest, the variances of the measurement equation for the output gap ($\hat{h}_{11}^2$) and the fiscal stance ($\hat{h}_{22}^2$) as well as their covariance ($\hat{h}_{12}^2$) for the benchmark TVP VAR, the regime-switching VAR and the constant parameter VAR. Moreover, the adjusted coefficients of determination, log-likelihoods, covariance matrix determinants and three information criteria as well as standardized measurement residuals are presented. Table 2 shows the same elements for alternative specifications of the TVP VAR.

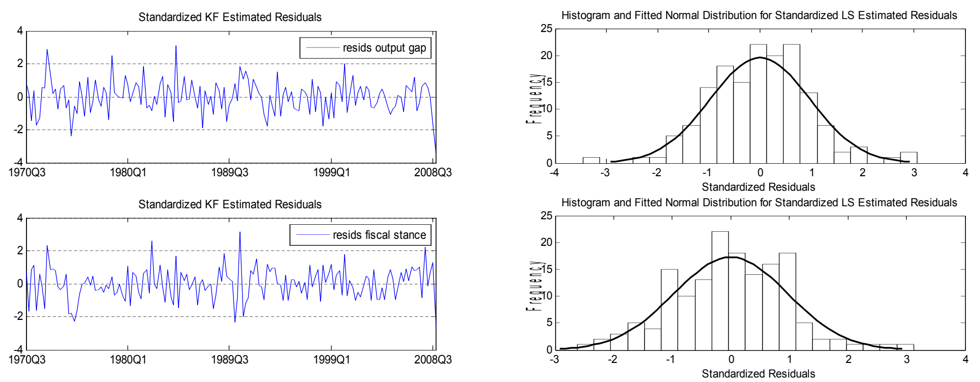
[3]We decided to use a direct search method called Nelder-Mead simplex search method, which is documented in Lagarias, Reeds, Wright, and Wright (1998). This method does not use numerical or analytic gradients, which makes sense because we want to keep the solution procedure simple and robust. In general, the chosen procedure can handle discontinuity, particularly if it does not occur near the solution. Alternatively, more efficient numerical methods such as gradient or scoring algorithms may be used. However, a scoring algorithm might have poorer convergence properties far from the optimum and has a high computational burden. Apart from the choice of an optimization algorithm, every procedure faces the problem that it might only converge to a local minimum.

Standardized measurement residuals

a) TVP VAR (ML Estimates, $\lambda = 10^{-9}$, Training Sample Priors)



b) Regime-Switching VAR (LS Estimates)



c) Benchmark VAR (LS Estimates)

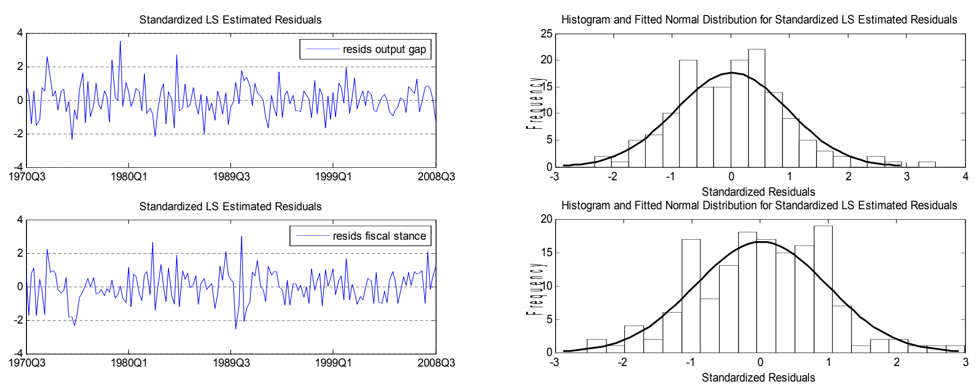
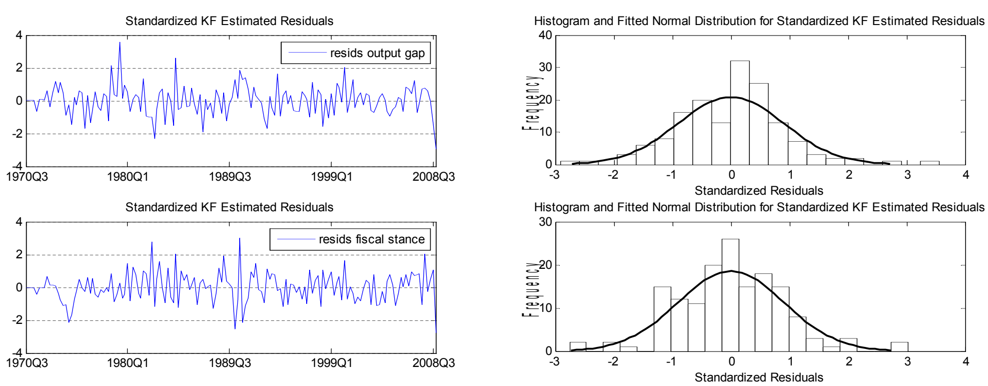


Table 2: Estimates of the Measurement Residual Variances Covariances and Model Statistics (Alternative TVP VAR Models) (Sample: 1970 Q3 – 2008 Q4)

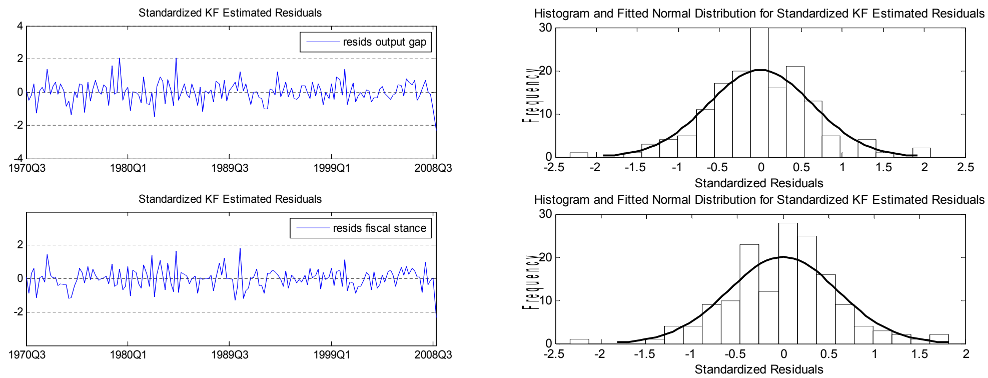
	d) TVP VAR	e) TVP VAR	f) TVP VAR	g) TVP VAR
Parameter and Model Statistic	ML Estimates, $\lambda = 10^{-9}$, Non-Informative Priors	ML Estimates, $\lambda = 10^{-5}$, Training Sample Priors	ML Estimates, $\lambda = 10^{-5}$, Non-Informative Priors	ML Estimates, Comprehensive Approach, Training Sample Priors
$\hat{h}_{11}^2$	$0.2753 * 10^{-4}$	$0.1282 * 10^{-4}$	$0.1258 * 10^{-4}$	$0.1614 * 10^{-4}$
$\hat{h}_{22}^2$	$0.2352 * 10^{-4}$	$0.1087 * 10^{-4}$	$0.1061 * 10^{-4}$	$0.0001 * 10^{-4}$
$\hat{h}_{12}^2$	$0.0589 * 10^{-4}$	$0.0595 * 10^{-4}$	$0.0578 * 10^{-4}$	$0.0046 * 10^{-4}$
$R^2_{adjusted}$(Eq. 1)	0.8614	0.9681	0.9703	≈1.0000
$R^2_{adjusted}$(Eq. 2)	0.9094	0.9815	0.9827	≈1.0000
Log Likelihood	1103.90	1166.80	1097.20	567.93
Determinant (Cov)	$0.6128 * 10^{-9}$	$0.1039 * 10^{-9}$	$0.1000 * 10^{-9}$	$0.1049 * 10^{-9}$
AIC	-21.110	-22.8849	-22.9231	-32.0860
SIC	-20.954	-22.7285	-22.7667	-31.9296

Standardized measurement residuals

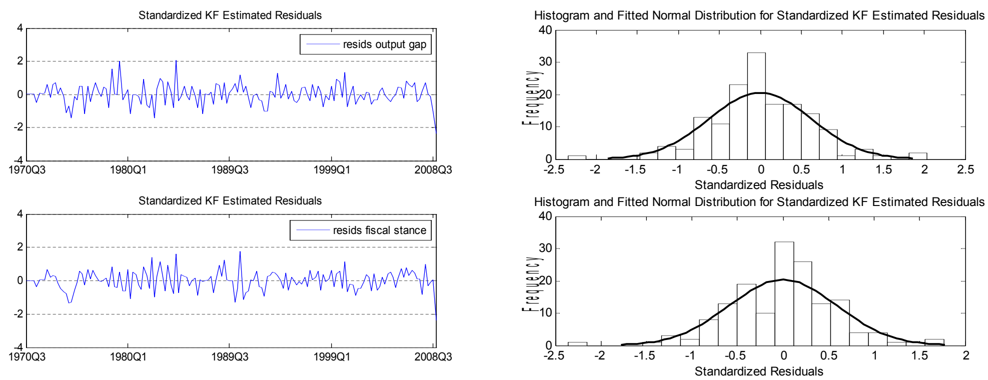
d) TVP VAR (ML KF Estimates, $\lambda = 10^{-9}$, Non-Informative Priors)



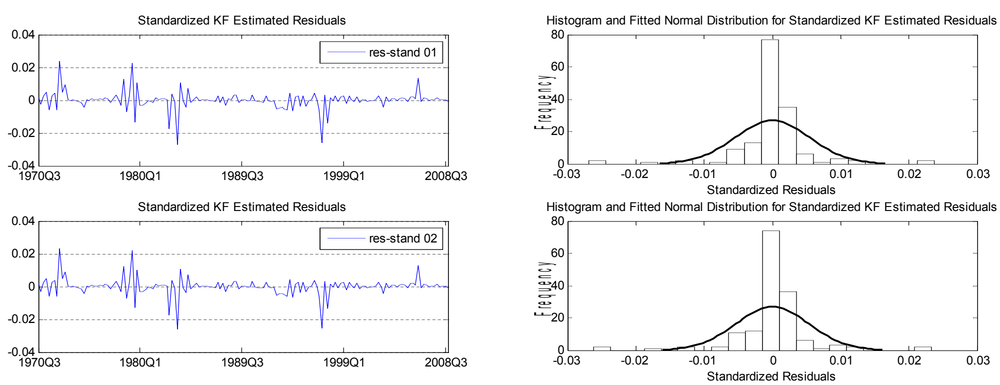
e) TVP VAR (ML KF Estimates, $\lambda = 10^{-5}$, Training Sample Priors)



f) TVP VAR (ML KF Estimates, $\lambda = 10^{-5}$, Non-Informative Priors)

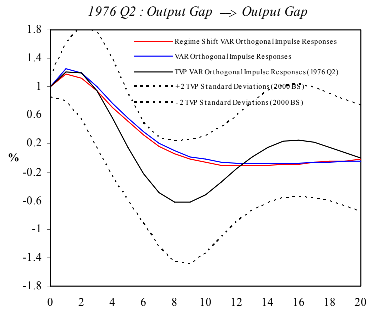


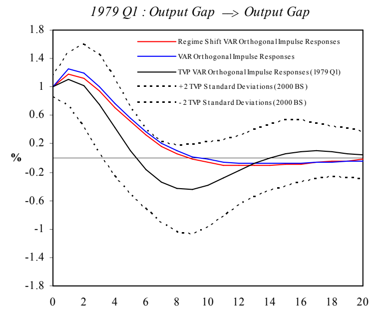
g) TVP VAR (ML KF Estimates, Comprehensive Estimation Approach (H and Q), Training Sample Priors)

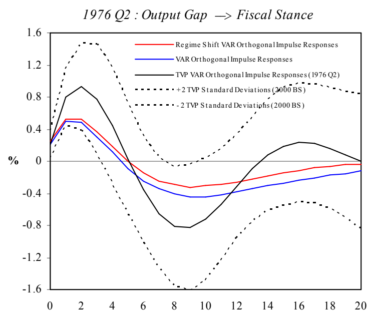


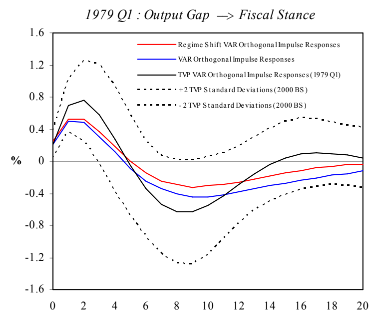
Appendix E: Impulse responses within bootstrapped standard deviations in the four identified regimes

Regime 1:

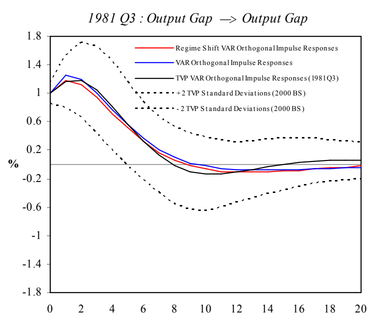


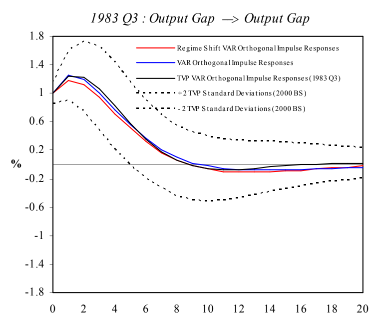


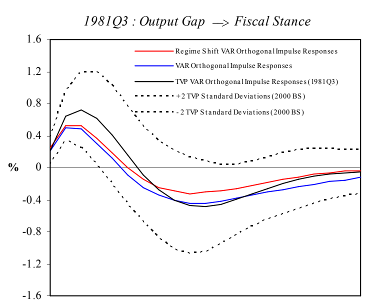


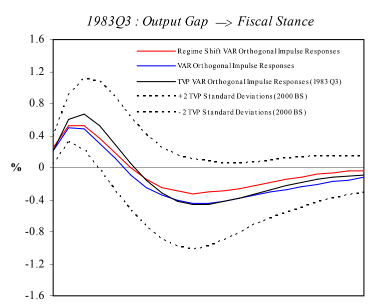


Regime 2:



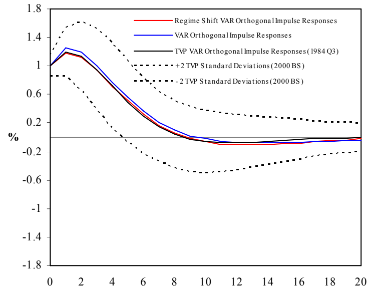




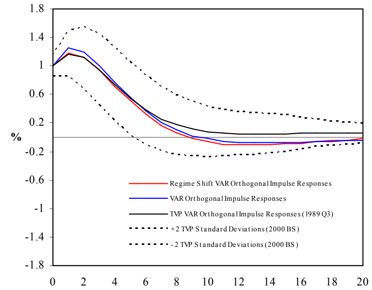


Regime 3:

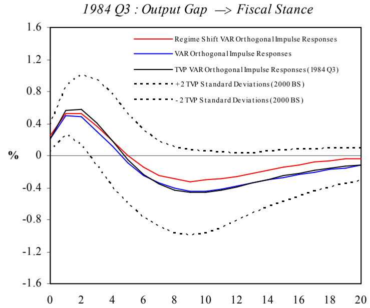
1984 Q3 : Output Gap —> Output Gap



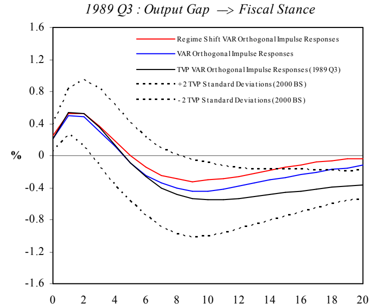
1989 Q3 : Output Gap —> Output Gap



1984 Q3 : Output Gap —> Fiscal Stance

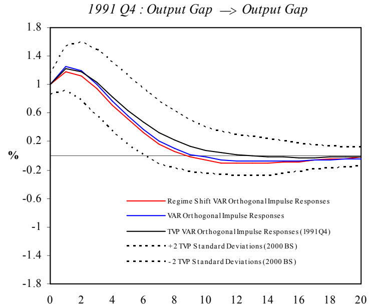


1989 Q3 : Output Gap —> Fiscal Stance

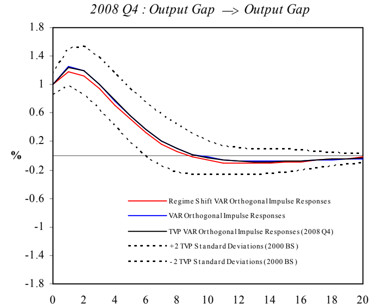


Regime 4:

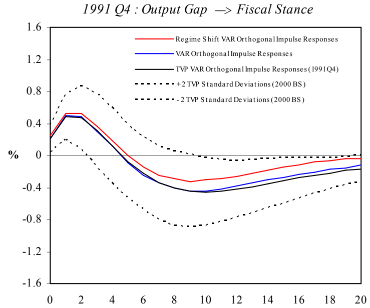
1991 Q4 : Output Gap —> Output Gap



2008 Q4 : Output Gap —> Output Gap



1991 Q4 : Output Gap —> Fiscal Stance



2008 Q4 : Output Gap —> Fiscal Stance