Nonlinearities in Swiss macroeconomic data

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1. INTRODUCTION

Since the seminal work of FRISCH (1933) and SLUTZKY (1937) business cycle models took the form of a linear impulse propagation mechanism. However, soon economists such as BURNS and MITCHELL (1935) as well as KEYNES (1936) pointed to possible asymmetries over the business cycle. Based on their observations different theoretical models such as the model with capacity constraints of HICKS (1950) or with downward rigid prices as in TOBIN (1972) evolved. However, only recently empirical methods were developed to test for the presence of asymmetries. With the surge of nonlinear methods business cycle asymmetries gained anew interest. While DELONG and SUMMER (1986) conclude that 'asymmetry is probably not a phenomenon of first importance' other researchers such as NEFTCI (1984) and SICHEL (1993) found significant evidence for the presence of asymmetry over the business cycle. The inconclusiveness of existing empirical tests spurred the search for alternative testing procedures and definitions of asymmetry. This paper will address mainly three questions: First, what is meant by asymmetry? Second, is there any evidence of asymmetry in Swiss macroeconomic data? Third, what are the possible causes of the observed nonlinearities?

The paper is organized as follows. Before testing for business cycle asymmetries in Switzerland, we have to give a brief account of theoretical models implying asymmetric reactions of the economy to monetary policy measures. The motivation for asymmetric effects of monetary policy and some underlying theoretical concepts are discussed in Section 2. A comparison of the presented theories then leads to a set of model-specific restrictions, which yield the basis of the empirical tests in the subsequent analysis. Since empirical evidence on Swiss data is very limited so far, I conduct a set of linearity tests on Swiss macroeconomic data in Section 3. Univariate tests as applied to 15 different Swiss time series reveal the presence of nonlinearity for a wide range of variables. In order to identify the possible causes and transmission channels of nonlinearity this analysis is extended to a multivariate test as suggested by GRANGER and TERÄSVIRTA (1993). The

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multivariate test results suggest that the effects of money on output are nonlinear depending on past values of money, prices and the exchange rate. Section 4 summarizes the main results and concludes the paper.

2. MONETARY THEORY AND ASYMMETRY

Since the definition of asymmetric effects of monetary policy depends decisively on the assumed transmission mechanism, it is necessary to distinguish between different models of the transition process. On one hand this helps to classify the possible forms of asymmetry in the effects of monetary policy. Model specific implications for the transmission mechanism on the other hand suggest a number of restrictions, which allow to test for alternative explanations of asymmetric effects of monetary policy in the empirical analysis.

In order to motivate the following empirical tests I distinguish between two classes of models: models in the Keynesian tradition, which are mainly based on nonlinear price rigidities, and models of imperfect capital markets, which stress information asymmetry as explanation for the effects of monetary policy on output and prices.

2.1 Models in the Keynesian tradition

Within the Keynesian tradition three different model classes have developed over the last 50 years: traditional Keynesian models with downward sticky prices, ‘Menu-cost’ models with up- and downward rigid prices and New Keynesian models with trend inflation. Traditional models explain asymmetric effects with asymmetric adjustments in prices and/or wages, e.g. downward rigid prices. The asymmetry of monetary policy depends thus on the sign of the respective change in money supply. The working hypothesis implied by traditional Keynesian models is thus:

\[ H_0 : \left| \frac{\partial y_t}{\partial c^-} \right| > \left| \frac{\partial y_t}{\partial c^+} \right| \quad \text{and} \quad \left| \frac{\partial p_t}{\partial c^-} \right| < \left| \frac{\partial p_t}{\partial c^+} \right| \]

with \( \left| \frac{\partial y_t}{\partial c^-} \right| \) representing the absolute effect of a negative monetary impulse on real output and \( \left| \frac{\partial p_t}{\partial c^-} \right| \) the respective absolute effect on prices.

However, missing microfoundation made these models less appealing. New models were developed explaining price stickiness with nominal rigidities. According to these so called ‘menu-cost’ models as e.g. in MANKIW (1985), firms are deterred from changing prices due to adjustment costs. Therefore, as long as adjustment costs exceed the expected increase in profits due to a price change, firms keep prices constant with the

1. See e.g. TOBIN (1972).
result that real variables adjust to the change in relative prices. Based on the latter models, I expect large monetary innovations to be neutral, triggering immediate price adjustments, while smaller innovations lead to an adjustment in real variables. Consequently, the working hypothesis of this approach can be summarized as:

\[ H_0 : \frac{\partial y_t}{\partial \epsilon_t^{\text{small}}} > \frac{\partial y_t}{\partial \epsilon_t^{\text{large}}} \quad \text{and} \quad \frac{\partial p_t}{\partial \epsilon_t^{\text{small}}} < \frac{\partial p_t}{\partial \epsilon_t^{\text{large}}} \]

However, this changes in an environment of non-zero expected inflation. Inspired by new models with state-dependent pricing (so called (S,s) approach) Ball and Mankiw (1994) demonstrate that size and sign of the monetary policy innovation is relevant for the expected effect on aggregate demand. Introducing expected positive inflation into a menu-cost model, Ball and Mankiw identify a combination of the two forms of asymmetry discussed above. The general price level is assumed to increase by a constant rate \( \pi \). The higher the rate of deterministic inflation, \( \pi \), and adjustment costs, \( D \), the stronger the asymmetry between positive and negative monetary policy shocks and consequently the nonlinearity between money and output. Thus depending on \( \pi \) and \( D \), increases in money supply will entail real effects only if they are very small, while non-neutrality might be observed even after a strong tightening. The working hypothesis of the state-dependent pricing model can thus be summarized as:

\[ H_0 : \frac{\partial y_t}{\partial \epsilon_t^{\text{small}}} > \frac{\partial y_t}{\partial \epsilon_t^{\text{large}}} \quad \text{and} \quad \frac{\partial p_t}{\partial \epsilon_t^{\text{small}}} < \frac{\partial p_t}{\partial \epsilon_t^{\text{large}}} \]

2.2 Models of imperfect capital markets

In contrast to the above mentioned New Keynesian models, explaining nonlinearity between money and output by introducing asymmetric price rigidities, models of imperfect capital markets stress information problems as relevant source of nonlinearity. Specifically, assuming financial markets to be incomplete, these models emphasize the credit channel as relevant transmission channel of monetary policy. The link between money

2. Models considering the effects of trend inflation in the presence of menu-costs can be found in Ball and Mankiw (1994), Caballero and Engel (1992) and Tsiddon (1993).

3. As pointed out in Blinder and Stiglitz (1983) also theories of credit rationing have to include some form of price rigidity, if demand shocks are to have real effects. Otherwise real variables would adjust quickly enough such that goods and credit markets are in equilibrium again.
and credit supply though is commonly assumed to change over the business cycle. There are two lines of reasoning why this should be so: monetary authorities are supposed to influence credit supply either directly via the 'bank lending channel' or indirectly via the 'balance sheet channel'.

For the bank lending channel the line of reasoning runs as follows. Generally, a tightening of monetary policy increases interest rates, which banks pass on to their customers and thus lead to a proportionate reduction in loan demand. If monetary authorities tighten when interest rates are already at a high level, a further increase in interest rates is assumed to increase default risks. As a consequence, an increase in loan contracts would lower expected profits of the bank. Banks therefore decide to turn down some credit requests, although clients are willing to pay the higher price of the loan contract. A reduction in liquidity supply during a period of strong growth by the Central Bank thus leads to an even stronger decline in borrowing than would be the case from higher interest rates alone. On the other hand, easing monetary policy in periods of sluggish growth would have no further impact as demand for credit is low anyway. Thus

\[ H_0 : \left| \frac{\partial y_{t,\text{expansion}}}{\partial \epsilon_t} \right| > \left| \frac{\partial y_{t,\text{recession}}}{\partial \epsilon_t} \right| \]

Almost the opposite conclusion is drawn based on the literature of the 'balance sheet channel'. According to these models individuals are much more sensitive to interest rate changes, when their balance sheets are weak. This is commonly assumed to be the case during troughs of the cycle as changes in the expected growth perspectives translate into an adjustment of asset values and thus in available collateral. In periods of high growth on the other hand, where balance sheets reflect a high degree of liquidity and collateral is easily provided, interest rate sensitivity is low and consequently monetary policy is assumed to have only minor effects on the credit market. The working hypothesis of models analyzing the balance sheet channel is then:

\[ H_0 : \left| \frac{\partial y_{t,\text{recession}}}{\partial \epsilon_t} \right| > \left| \frac{\partial y_{t,\text{expansion}}}{\partial \epsilon_t} \right| . \]

4. For extended discussion of the relationship between monetary policy and credit supply see e.g. Bernanke and Gertler (1989), Blinder and Stiglitz (1983) and Stiglitz and Weiss (1992).
5. See Bernanke and Gertler (1995) and de Bondt (1998) for further discussion of the two channels.
6. For an extensive analysis of moral hazard and/or adverse selection in the credit market see Stiglitz and Weiss (1992). A general survey on credit rationing can be found in Jaffee and Stiglitz (1990) and Freixas and Rochet (1997).
The difference between these two model classes can be easily demonstrated in Graph 1 and Graph 2. Both graphs depict the consequences of a positive and a negative aggregate demand shock in an aggregate demand-aggregate supply model.\footnote{This graphical illustration of differences between both model classes is taken from KARRAS (1996).}

**Graph 1: Asymmetric price rigidity in Keynesian models**

As models in the Keynesian tradition imply some form of nonlinear price adjustment, the aggregate supply curve is assumed to be convex. Consequently, an increase in aggregate demand will lead to a stronger reaction in prices than a reduction in demand, i.e. $|p_1 - p_0| > |p_2 - p_0|$. The contrary is true in models based on imperfect capital markets, which do not request a nonlinear rigidity in prices. Consequently, the supply curve in these models is linear as illustrated in Graph 2. In this case the asymmetry is ‘demand driven’, i.e. asymmetry in prices is positively correlated with asymmetry in output such that $|p_2 - p_0| > |p_1 - p_0|$.

Therefore the comparison of asymmetry in output and prices constitutes a test for the two competing causes of asymmetric effects of monetary policy.
After the motivation of asymmetric effects has been discussed, one might raise the question whether these theories have any relevance for the business cycle dynamics in Switzerland. Allowing for asymmetries in empirical models will seem reasonable only, if business cycle asymmetries represent a significant characteristic in the behavior of actual data. Therefore I conduct a series of linearity tests on Swiss macroeconomic time series which provide information about the possible form and source of nonlinearity observed.

3. TESTING FOR BUSINESS CYCLE ASYMMETRY IN SWISS DATA

This section will address mainly two questions:

(1) Is there any evidence of asymmetry in Swiss macroeconomic data?
(2) What are the possible causes of the observed nonlinearities?

Apparently, little research has so far been devoted to potential asymmetric behavior of Swiss macroeconomic time series. DANTHINE and NEFTCI (1990) measured the form of the Swiss business cycle with the help of a shape function. Comparing these results for Swiss data with the same analysis of German data, DANTHINE and NEFTCI observed that ‘[…] the upturns in Germany are much sharper than the upturns in Switzerland. It
appears that Swiss economy recovers from Business Cycles in a gradual fashion. This in-
fers that the Swiss business cycle resembles more the traditional Keynesian proposition
of sharp downturns and slow recoveries. Two further papers studied asymmetric adjust-
ments in Swiss output growth, focusing though on a specific parametrization of nonli-
nearity. While Lenz (1997) tested for asymmetric effects of monetary policy on real
GDP, Lüscher (1998) modeled a nonlinear Phillips curve for the dynamic relation be-
tween real growth and inflation. All three papers on business cycle asymmetries in Swit-
zerland found evidence for their respective form of asymmetry. However, all of these
studies lacked an initial test for the presence of nonlinearity before estimating the speci-
fic nonlinear model. Knowing about the danger of spuriously good fits of nonlinear
models, testing for linearity becomes a conditio sine qua non before estimating a non-
linear model and is thus performed below.9

The following tests are based on the concept of business cycle asymmetry as suggested
in Potter (1995). According to this concept nonlinear models can give rise to three dif-
ferent forms of asymmetry. Consequently, testing for the presence of nonlinearity yields
a general test for asymmetry. The following section introduces the parametric test for
nonlinearity used in the analysis of section 3.2 and 3.3.

3.1 A parametric linearity test

The parametric linearity test applied in this section serves as a test for the presence of
asymmetry. The presentation of the test procedure reveals that the latter can also be
used to examine the underlying form of nonlinearity. Therefore, addressing question (i)
and (ii) of this section, I use the LM test introduced by Luukkonen et al. (1988) in or-
der to examine the possible form and cause of asymmetric behavior of Swiss macrodata.
This test procedure has been chosen for two reasons. First, the power of the LM test was
found to be superior compared to other tests.10 Second, since the test represents a test

10. Simulating 1000 replications from a logistic smooth transition autoregressive model (LSTAR)
with a sample size of 100 and 300 observations, the empirical size of the overall linearity test was
found to be about 3-4% (compare Teräsvirta (1994)). In this test, linearity was rejected if the p-
value of the test was below 0.01. However, since the underlying form of nonlinearity is unknown,
one might prefer testing linearity against an unspecified alternative such as the RESET test, the
CUSUM test or a neural network test instead (see Granger and Teräsvirta et al. (1993) for
detailed discussion of these tests). However, according to several Monte Carlo studies, the LM
test according to Luukkonen et al. (1988) was found to be more powerful than these tests
against an unspecified form of nonlinearity (compare results in Teräsvirta et al. (1993) and
Luukkonen et al. (1988)). Since this was also true (at least for specific forms of the LM test), if
the data generating process followed another form of nonlinearity, the test might also have power
against alternative nonlinear models. However, the specific version of the test applied in this
study (referred to as LSTAR4 in Teräsvirta et al. (1993)) was found to have relatively low
power against linearity, if the underlying data generating process was e.g. a neural network
model, which might therefore reduce the probability of rejecting linearity against other forms of
nonlinearity.
against a specific form of nonlinearity, a rejection of linearity infers a specific form of asymmetry. However, it should be pointed out that this test has also power against other forms of nonlinearity. Therefore, inference on the specific form of nonlinearity should be considered as preliminary until specific nonlinear models are estimated.

Due to the advantages pointed out above, the LM test procedure proposed by Luukkonen et al. (1988) is applied hereafter. In order to keep the form of nonlinearity as general as possible, Luukkonen et al. suggest to parameterize nonlinearity in the form of a Smooth Transition Autoregressive model (STAR). The STAR model can be represented as follows:

$$y_t = \theta_1' w_t + \theta_2' w_t F(y_{t-d}) + u_t$$ (1)

where $$\theta_j = (\theta_{j1}, ..., \theta_{jp})'$$, $$w_t = (1, y_{t-1}, ..., y_{t-p})'$$ and $$u_t \sim i.i.d.(0, \sigma^2)$$. p represents the order of the AR process. $$y_{t-d}$$ is the transition variable with lag d, which introduces a specific form of nonlinearity. Depending on the form of the transition function $$F()$$, which determines the form of nonlinearity inherent, one distinguishes between logistic transition models (LSTAR),

$$F(y_{t-d}) = (1 + exp[-\gamma(y_{t-d} - c)])^{-1}, \quad (2)$$

and exponential transition models (ESTAR),

$$F(y_{t-d}) = 1 - exp[-\gamma(y_{t-d} - c)^2]. \quad (3)$$

An example for these two specifications of the STAR model is given in Graph 3 and Graph 4 respectively.

A comparison of these two graphs reveals that the nonlinearity captured in the specific form of the transition function $$F()$$ encompasses asymmetry in size or sign. Take e.g. the LSTAR model illustrated in Graph 3. As the transition function takes on a different value when $$y_{t-d}$$ is above or below the threshold c (e.g. $$F(-4) \neq F(4)$$), asymmetry in sign should consequently manifest itself in a logistic form of the STAR model. The LSTAR model might therefore imply different behavior in recessions and expansions. Instead, the ESTAR model depicted in Graph 4 depicts a form of asymmetry with respect to the absolute difference between the transmission variable $$y_{t-d}$$ and the threshold c and could therefore represent an asymmetry in size.

The test of linearity against a STAR model can be formulated as a LM test. This is especially appealing since the nonlinear model does not have to be estimated. An auxiliary regression on the estimated residuals under the Null can be used in order to test for linearity instead.11

Graph 3: $F()$ of logistic STAR model

Specification as given in equation (2) with $\gamma = 1$ and $c = 0.5$

Graph 4: $F()$ of exponential STAR model

Specification as given in equation (3) with $\gamma = 1$ and $c = 0.5$
\[ \hat{u}_t = c + \sum_{j=1}^{p} \beta_{1j} y_{t-j} + \sum_{j=1}^{p} \beta_{2j} y_{t-j} y_{t-d} + \sum_{j=1}^{p} \beta_{3j} y_{t-j}^2 y_{t-d} + \sum_{j=1}^{p} \beta_{4j} y_{t-j}^3. \]  

(4)

Depending on the restrictions imposed on the \( \beta \)'s, the form of nonlinearity will either represent an asymmetry in size, sign or history.

In order to test the different hypotheses, the LM statistic is computed, with \( \text{LM} = T(\text{SSR}_0 - \text{SSR}_1)/\text{SSR}_0 \), where \( T \) is the number of observations, \( \text{SSR}_0 \) and \( \text{SSR}_1 \) the sums of squared residuals under \( H_0 \) and \( H_1 \) respectively. The LM statistic has a \( \chi^2(3p) \) distribution. For small samples Teräsvirta (1994) suggests using the equivalent F statistic instead, where \( \text{F} = ([\text{SSR}_0 - \text{SSR}_1]/3p)/[\text{SSR}_0/(T - (4p + 1))] \).

Searching for a possible source of asymmetry, the univariate LM test has to be extended to the multivariate version of the LM test as proposed in Granger and Teräsvirta (1993). In the multivariate version any lagged variable can serve as transition variable. Consequently, this test should yield an answer to what might be the possible cause of the observed nonlinearities. The model can be represented in the following form:

\[ y_t = \theta'_1 w_t + \theta'_2 w_t F(s_{t-d}) + u_t \]

where \( w_t = (1, y_{t-1}, ..., y_{t-p}, x_{t,t}, ..., x_{k,t})' \), \( s_{t-d} \) represents any lagged endogenous and/or any exogenous variable with \( E(s_{t-d} u_t) = 0 \), \( \theta'_j = (\theta_{2j}, \theta_{3j}, ..., \theta_{mj})' \) where \( m = p + k \) and \( u_t \sim \text{i.i.d.}(0, \sigma^2) \) and \( p \) the order of the AR process. The number of nonlinear relations in this model is determined by the coefficient vector \( \theta_2 \). If \( \theta_2 \) is unrestricted, i.e. \( \theta_{2,i} \neq 0 \) for all \( i \), asymmetry is allowed to disperse via all endogenous variables included in the vector \( w_t \). However, if \( \theta_2 \) is restricted such that \( \theta_{2,j} \neq 0 \) and \( \theta_{2,i} = 0 \) for all \( i \neq j \), the endogenous \( y_t \) may react nonlinearly only with respect to \( w_j \).

For the univariate case the test procedure can be summarized in the following 4 steps:

1. Step: Test for linearity
\( H_{0,1}: \beta_2 = \beta_3 = \beta_4 = 0 \) with \( \beta_i \) being a vector of dimension \([p,1]\)

2. Step: Test for ESTAR against LSTAR model
\( H_{0,2}: \beta_4 = 0 \) against \( H_{1,2}: \beta_4 \neq 0 \)
A rejection of \( H_{0,2} \) can be interpreted as a rejection of the ESTAR model. Note however, that in the special case when the LSTAR model takes the form where only the unconditional mean of the time series changes between regimes, this test might not lead to a conclusive answer. Thus one may also test between LSTAR against ESTAR model following step 3 and 4:

3. Step:
\( H_{0,3}: \beta_3 = 0 \mid \beta_4 = 0 \) against \( H_{1,3}: \beta_3 \neq 0 \mid \beta_4 = 0 \)

4. Step:
\( H_{0,4}: \beta_2 = 0 \mid \beta_3 = \beta_4 = 0 \) against \( H_{1,4}: \beta_2 \neq 0 \mid \beta_3 = \beta_4 = 0 \)
Rejecting step 4 after having accepted the test in step 3, supports the choice of a logistic form of $F$. In the reverse case, the dynamics are better represented by an ESTAR model.

### 3.2 Univariate linearity tests for Swiss data

Trying to explore the relevance of asymmetric behavior in Swiss data, I study the behavior of different Swiss macroeconomic variables over the business cycle. In the following section the univariate LM test is conducted on 15 Swiss macroeconomic time series over the sample period 1975:1 to 1997:4. However, in order to guarantee a standard distribution of the test, the time series are required to be stationary. Therefore, the cyclical component of the level data has to be extracted by detrending the data. Applying a non-linear filter though might introduce asymmetry to the data, such that the test results become a mere statistical artifact of the detrending method used.\(^\text{12}\) Consequently, different detrending methods are applied on logged variables: linear detrending, the HP filter and first differences. The choice of applied detrending methods is based on the following consideration. While the latter two filters have a linear representation and are therefore unlikely to induce additional asymmetry, linear detrending fails to fulfill this requirement. Despite the possibility of adding asymmetry by linear detrending, I choose this method as an alternative filter since several authors report the rejection of linearity solely due to structural breaks.\(^\text{13}\) I therefore account for structural breaks in linear detrending, when indicated by the procedure of PERRON (1997) for unknown structural breaks. Due to the trade-off between avoiding additional asymmetry – by using a linear filter – and accounting for structural breaks, it is necessary to compare the results of linear detrended data to the findings based on linearly filtered data. The HP filter and first differences instead fulfill stationarity as well as linear representation. Accordingly, these filters can be used for comparison to linear detrended data.\(^\text{14}\)

\(^{12}\) Note, that according to CANOVA (1998), a wide class of detrending methods induce asymmetry, observing significant skewness in the detrended time series. However, testing the characteristics of detrended data against normality, as done by CANOVA, seems insufficient to attribute observed skewness to the specific method of detrending used. Instead the cyclical component itself might be asymmetric.

\(^{13}\) Compare RAVN and SOLA (1996), TERÄSVIRTA and ANDERSON (1992) and WEISE (1999).

\(^{14}\) See LUUKKONEN and TERÄSVIRTA (1991) for the invariance of test results to data transformation.
Table 1: F-statistics of univariate linearity tests

Regression:  \[ u_t = \beta_1 w_t + \beta_2 w_t y_{t-d} + \beta_3 w_t y_{t-d}^2 + \beta_4 w_t y_{t-d}^3 + \epsilon_t \]

Test:  \[ H_0 : \beta_2 = \beta_3 = \beta_4 = 0 \quad H_1 : \beta_2 = \beta_3 = \beta_4 \neq 0 \]

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<th>Variable</th>
<th>Linear detrended</th>
<th>HP filtered</th>
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<td>2.59 **</td>
<td>5.82 **</td>
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<td>(0.28)</td>
<td>(0.02)</td>
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<td>2.53 **</td>
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<td></td>
<td>(0.05)</td>
<td>(0.03)</td>
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<td>1.31</td>
<td>1.98 **</td>
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<td></td>
<td>(0.19)</td>
<td>(0.22)</td>
<td>(0.03)</td>
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<td>ind. production</td>
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<td>11.90 **</td>
<td>12.13 **</td>
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<td>3.32 **</td>
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<td>(0.00)</td>
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<td>(0.09)</td>
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<td></td>
<td>(0.04)</td>
<td>(0.01)</td>
<td>(0.14)</td>
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Note: The sample is quarterly data in logs from 1975:1–1997:4. However, interest rates as well as the term spread are stationary over the sample period. Level data is used for these three time series. Based on the tests suggested by Perron (1997) the following dates for structural breaks are chosen for linear detrending the following variables in logs: GDP: 77:4, investment: 86:4, net exports: 77:4, ind. production: 89:3, CPI: 89:2, WPI: 85:4, M1: 79:1, M2: 78:2, Credit: 88:2, nom. exchange rate: 78:1, real exchange rate: 78:3. The usual lambda for quarterly data of 1600 is used in case of the HP filter. For data description see Appendix A.

The test is based on the auxiliary regression presented in the table, where \( w_t = (1, y_{t-1}, \ldots, y_{t-p})' \). \( p \) is chosen according to the Ljung-Box Q-statistic as to eliminate observed serial correlation. \( d \) varies between 1 and 5 and is determined by the significance level of the test statistic. Significance levels are given in parentheses. ** and * indicate significance at 5% and 10% levels respectively.
As a first step, the number of included lags, \( p \), is chosen according to the Akaike criterion. However, using the Akaike criterion, the Ljung-Box Q-statistic still indicates significant serial correlation in the error terms. As pointed out by Teräsvirta (1994), serial correlation might bias the test towards rejecting the Null, i.e. rejecting linearity. Therefore I use a more conservative approach, augmenting the number of lags until serial correlation is no longer present. As a second step, the transition variable \( y_{t-d} \) has to be determined. Following the suggestion of Granger and Teräsvirta (1993), I test for lags \( d \) between 1 and 5 and choose the lag, where linearity is rejected with highest significance, to represent the transition variable. Table 1 reports only the results for lag \( d \) with the highest significance.

In numerous cases, linearity can be rejected commonly for all detrending methods. This is especially true for the first two methods. The similar test results of linear detrended and HP filtered data infer no additional nonlinearity is added by the former detrending method despite its lack of a linear representation. The time series for which linearity can be rejected across all detrending methods are real investment, industrial production, wpi and ml. For linear detrending and HP filtered data, I reject linearity furthermore in the case of credit, 3-Month Euro rates, the term spread as well as for both exchange rate variables.

While the univariate linearity test indicates the presence of nonlinearity, it does not yield any answer to what causes the asymmetry observed. Therefore I extend the test procedure to a multivariate case in the following section.

### 3.3 Multivariate linearity tests for Swiss data

Since the sample under consideration (1975:1–1997:4) is rather short, the system should be restricted to a relatively small number of coefficients. Following the traditional VAR-literature, I choose a system including logged data for real GDP \( (y_t) \), the Wholesale Price Index \( (p_t) \) and M1 \( (m_t) \).\(^{15}\) As Switzerland is a very open economy the trade-weighted nominal exchange rate \( (e_t) \) is added as a fourth variable, such that \( w_t = (1, y_{t-1}, \ldots, y_{t-p}, p_{t-1}, \ldots, m_{t-1}, \ldots, e_{t-p})' \). I use linear detrended data for the subsequent tests. This decision is based on the following consideration. As I am testing for business cycles asymmetries, the detrending method applied should emphasize lower frequency ranges. With first differencing stressing mainly high frequencies, the HP filter and linear detrending seem more appropriate since both filters emphasize lower frequency ranges. Furthermore, as the results of the univariate parametric test differed only slightly between HP and linear detrended data, results based on these two filters are expected to yield similar results also in the multivariate case. Since the univariate

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15. The wholesale price index was used instead of the consumer price index for two reasons. First, the Swiss CPI is known to exhibit especially high persistence due to long lags in rent adjustments (See Etting (1995)). Furthermore, this time series was redefined several times during the sample period making it thus subject to structural breaks.
tests above did not identify a distinct source of nonlinearity, I start by allowing the influence of all explanatory variables to vary across different regimes. In this case \( \theta_2 \) in equation (5) remains unrestricted.

Table 2: Multivariate linearity tests with unrestricted transmission channel

<table>
<thead>
<tr>
<th>Regression: ( \hat{u}<em>t = \beta_1 w_t + \beta_2 w_t y</em>{t-d} + \beta_3 w_t m_{t-d} + \beta_4 w_t e_{t-d} + \epsilon_t )</th>
</tr>
</thead>
<tbody>
<tr>
<td>Test: ( H_0 : \beta_2 = \beta_3 = \beta_4 = 0, H_1 : \beta_2 = \beta_3 = \beta_4 \neq 0 )</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Transition variable ( s_{t-d} )</th>
<th>( y_{t-d} )</th>
<th>( m_{t-d} )</th>
<th>( p_{t-d} )</th>
<th>( \epsilon_{t-d} )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( w_t )</td>
<td>5.48 **</td>
<td>2.90 **</td>
<td>3.69 **</td>
<td>4.69 **</td>
</tr>
<tr>
<td></td>
<td>(0.00)</td>
<td>(0.03)</td>
<td>(0.01)</td>
<td>(0.00)</td>
</tr>
<tr>
<td>( w_t )</td>
<td>3.66 **</td>
<td>4.19 **</td>
<td>2.55 **</td>
<td>2.94 **</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.00)</td>
<td>(0.05)</td>
<td>(0.03)</td>
</tr>
<tr>
<td>( w_t )</td>
<td>1.54</td>
<td>1.33</td>
<td>1.92</td>
<td>1.93</td>
</tr>
<tr>
<td></td>
<td>(0.20)</td>
<td>(0.27)</td>
<td>(0.12)</td>
<td>(0.12)</td>
</tr>
<tr>
<td>( w_t )</td>
<td>3.15 **</td>
<td>2.83 **</td>
<td>2.53 **</td>
<td>2.25 *</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.05)</td>
<td>(0.07)</td>
</tr>
</tbody>
</table>

Note: The sample is quarterly logged data 75:1-97:4. Variables are linearly detrended with structural breaks as indicated in Table 1. For data description see also Appendix A. This test is based on the auxiliary regression presented in the table, where \( w_t = (1, y_{t-1}, \ldots, y_{t-p}, \ldots, \epsilon_{t-p})' \). \( p = 4 \), such that according to the Ljung-Box Q-statistic serial correlation is no longer present. \( d \) varies between 1 and 5 and is determined by the significance level of the test statistic. Significance levels are given in parentheses. ** and * indicate significance at 5 % and 10 % levels respectively.

The regimes are either determined by the state of \( y_{t-d}, m_{t-d}, p_{t-d}, \) or \( \epsilon_{t-d} \), chosen as the relevant transition variable \( s_{t-d} \). I follow the same procedure as above for the choice of the number of lags \( p \) and the lag of the transition variable \( d \). Again, only the results of lag \( d \) for which the test statistic reaches the lowest p-value are reported in Table 2. In order to keep the illustration traceable, I renounce on reporting the results of the tests for the specific form of nonlinearity. Instead the results of these tests are discussed in the text, where the former clearly point to a specific from of the transition function \( F \).

Prices exhibit no asymmetry with respect to any transition variable. For all other variables instead, it is difficult to identify the relevant transmission variable. Irrespective of the variable used as transition variable, linearity can be rejected. Accordingly, output, money and the exchange rate react nonlinearly to changes in either of the explanatory variables in \( w_t \) depending on the particular state of the transition variable. The results suggest an omnipresence of asymmetry in contrast to WEISE (1999), who could not reject linearity in output and prices in U.S. data when using money as transition variable. With the theoretical models in mind, the test results might be interpreted in the following way. As presented in Section 2, traditional Keynesian models suggest that downward rigid
prices give rise to asymmetric effects of monetary policy. This class of models thus implies a nonlinearity in output and prices depending on the monetary regime. The significance of lagged money as the transition variable in Table 2 can be seen as first evidence for the traditional Keynesian asymmetry. However, not observing any evidence of nonlinearity in prices when money is used as transition variable, is at odds with this theory, as these models rely on asymmetric price adjustments. On the other hand, models explaining asymmetric effects of money on output via credit constraints seem to find more support in the empirical results. Models of imperfect capital markets assume that the effectiveness of monetary policy changes depending on the state of the cycle. Consequently, the effect of money on output and prices should depend on the state of the business cycle. Output is thus expected to be the relevant transition variable. Results in Table 2 indicate nonlinearity in output depending on lagged output as transition variable. However, since linearity in output is rejected significantly with respect to all transition variables, these tests do not allow to discriminate among the latter.

Therefore in order to pin down the most relevant cause of asymmetry in money, one might compare different transmission channels of nonlinearity. This is achieved by restricting $\theta_2$ to represent a Null vector except for the coefficients on the relevant transmission variables in $\tilde{w}_t$. Restricting the dispersion of nonlinearity in money e.g. on lags 1 through $p$ of the exchange rate, implies a restricted coefficient vector $\tilde{\theta}_2' = (0, \ldots, 0, \theta_2, m_{-p}, \theta_2, m_{-p-1}, \ldots, \theta_2, m)$. Table 3 reports the test results for different transmission channels.

**Table 3: Multivariate tests with money as relevant transition variable and restricted transition channel**

| Regression: $\hat{u}_t = \beta_1 w_t + \beta_2 \tilde{w}_t m_{t-d} + \beta_3 \tilde{w}_t m_{t-d}^2 + \beta_4 \tilde{w}_t m_{t-d}^3 + \epsilon_t$ |
|---|---|---|---|---|
| Test: $H_0: \beta_2 = \beta_3 = \beta_4 = 0, H_1: \beta_2 = \beta_3 = \beta_4 \neq 0$ |
| transition variable | $m_{t-d}$ | $m_{t-d}$ | $m_{t-d}$ | $m_{t-d}$ |
| transmission channel $\tilde{w}_t$ | $y_{t-j}$ | $m_{t-j}$ | $p_{t-j}$ | $e_{t-j}$ |
| dependent |
| $y_t$ | 0.96 | 1.16 | 2.03 * | 1.90 |
| | (0.44) | (0.34) | (0.10) | (0.12) |
| $m_t$ | 3.21 ** | 1.86 | 1.93 | 4.47 ** |
| | (0.02) | (0.13) | (0.12) | (0.00) |
| $p_t$ | 1.36 | 0.95 | 1.06 | 0.96 |
| | (0.26) | (0.44) | (0.39) | (0.44) |
| $e_t$ | 3.53 ** | 0.97 | 0.77 | 2.82 ** |
| | (0.01) | (0.43) | (0.55) | (0.03) |

Note: The sample is quarterly logged data 75:1-97:4. This test is based on the auxiliary regression presented in the table, where $\tilde{w}_t = (x_{t-1}, \ldots, x_{t-p})'$, where $x_t = y_t, p_t, m_t$ and $e_t$ respectively and $j = 1$ to $p$. Compare also footnote to Table 2. Significance levels are given in parentheses. ** and * indicate significance at 5% and 10% levels respectively.
Note that prices represent the only significant variable, which affects output differently depending on the monetary regime. Again, the test results do not indicate any nonlinear reaction in prices, where the tests fail to reject linearity even at a 10% level. Although any of the above presented models is based on some form of price rigidity, and thus some nonlinearity in price adjustments with respect to changes in money, this phenomenon does not show up in Table 3.

As pointed out above, I extend the analysis of the univariate test further by testing for the specific form of nonlinearity. For all dependent variables and each respective transmission channel in Table 3 I conduct the tests of step 2 to 4. The test results indicate an exponential form of nonlinearity for the following relations: (1) for the effect of prices on output, (2) for the effects of output on money and (3) for the reaction of the exchange rate on its own past development. Instead, a LSTAR model should be chosen (1) for the dynamic adjustment in money to changes in the exchange rate as well as (2) in the effects of output changes on the exchange rate.

What conclusions can be drawn based on the test results of step 2 to 4? As suggested in Graph 4, a possible interpretation of the exponential form is that the dependent variable reacts more or less strongly depending on the absolute distance of lagged money from its specific threshold value c, i.e., an asymmetry in size. I interpret the exponential form of nonlinearity in output with respect to prices as to suggest the presence of price rigidity according to a menu-cost model. Analogously, money seems to react nonlinearly to changes in output. This infers a stronger reaction of money to real growth depending on how much money diverges from some threshold c.

The interpretation of the test results differ if the indicated nonlinear model takes the form of an LSTAR model. Graph 3 has illustrated that this form of nonlinearity can be interpreted as an asymmetry with respect to sign. Evidently, the LSTAR model for the relation between money and the exchange rate suggests a different response in money supply and/or money demand to exchange rate changes depending on the sign of lagged changes in money. The latter might be translated as restrictive or expansionary monetary conditions. The effects of changes in real income on the exchange rate seem to be characterized by the same asymmetry. In analogy to the logistic form of nonlinearity in money, the exchange rate might thus behave differently depending on whether past changes in money were positive or negative.

Consequently, I find some evidence that prices are a source of asymmetric dynamics in Swiss macrodata. However, no asymmetric price adjustment is observed independently of the transmission channel used. At this point, one might object that this test procedure does not use unanticipated monetary shocks as transition variable as normally used in the literature of asymmetric effects of monetary policy. However, this form of nonlinearity should also be observable in responses to the systematic part of monetary policy, if adjustment costs are the reason for price rigidity. The strong presence of nonlinearity in money and the exchange rate hints to another important source

16. A detailed description of the test procedure can be found in Teräsvirta (1994).
of nonlinearity, that received little attention in the literature on business cycle asymmetry so far.

According to the models attributing nonlinear effects of monetary policy to credit constraints, the effectiveness of monetary policy depends on the state of the business cycle. This form of nonlinearity can be represented in a model, where output is used as transition variable. In order to test whether the effectiveness of money depends on the state of any explanatory variable, I proceed by constraining the transmission channel of nonlinearity on money. This restriction implies different effects of money on the four variables depending on the state of output, prices, money and the exchange rate. Test results are reported in Table 4.

In general, linearity can not be rejected by any of the tests. The only exception is the nonlinearity in money depending on past price and exchange rate developments, both best characterized by an exponential form of the STAR model. Since the earlier tests found no evidence for asymmetric price adjustment as predicted by traditional and New Keynesian models, credit constraints might have been expected to be the primary source of a nonlinear relation between money and output. However, linearity in output can not be rejected for any transition variable, if the transmission channel is restricted on money only. The same is true for prices and the exchange rate.

Table 4: Multivariate tests with monetary transmission channel

| Regression: | $\hat{y}_t = \beta_1 w_t + \beta_2 \bar{w}_t s_{t-d} + \beta_3 \bar{w}_t p_{t-d} + \beta_4 \bar{w}_t e_{t-d} + \epsilon_t$ |
| Test: | $\beta_2 = \beta_3 = \beta_4 = 0$ |
| transition variable $s_{t-d}$ | $y_{t-d}$ | $m_{t-d}$ | $p_{t-d}$ | $e_{t-d}$ |
| transmission channel $\bar{w}_t$ | $m_{t-j}$ | $m_{t-j}$ | $m_{t-j}$ | $m_{t-j}$ |
| dependent | | | | |
| $y_t$ | 1.25 | 1.16 | 1.67 | 1.40 |
| | (0.30) | (0.34) | (0.17) | (0.24) |
| $m_t$ | 1.72 | 1.86 | 2.95 ** | 2.84 ** |
| | (0.16) | (0.13) | (0.03) | (0.03) |
| $p_t$ | 1.22 | 0.95 | 1.84 | 1.31 |
| | (0.31) | (0.44) | (0.13) | (0.28) |
| $e_t$ | 1.73 | 0.97 | 1.82 | 1.41 |
| | (0.16) | (0.43) | (0.14) | (0.24) |

Note: The sample is quarterly logged data 75:1-97:4. This test is based on the auxiliary regression presented in the table, where $\bar{w}_t = (m_{t-1}, ..., m_{t-p})^j$, $j = 1$ to $p$. Compare also footnote to Table 2. Significance levels are given in parentheses. ** and * indicate significance at 5% and 10% levels respectively.

In order to evaluate the robustness of the tests, I repeat the tests on a system using industrial production instead of real GDP as well as the real instead of the nominal exchange rate. Tables for these test results are omitted here, but are available on request. Tests for money as the relevant transition variable remain almost the same, with the exception
that money is now the relevant transmission channel of nonlinearity in output. In contrast to Table 3 though, this does not affect the dynamics between output and prices.

When money is assumed to be the only transmission channel of asymmetry, linearity can be rejected in industrial production for the transition variables money and prices at a 7% and 2% significance level respectively. According to the respective test results for step 2 to 4 the logistic form of nonlinearity seems the best choice in both cases. The interpretation of the LSTAR model implies that the effects of money on industrial production depend on whether monetary conditions have been restrictive or expansionary.

In general, the data does not support the dependence on the business cycle stage as proposed by models with credit constraints. Consequently, I can draw two possible conclusions. Either lagged values of output fail to indicate the presence of credit constraints and another exogenous variables, as e.g. the net financial wealth of firms, should be used instead, or credit constraints might not be the primary source of nonlinearity between money and output.

Using the real instead of the nominal trade weighted exchange rate changes the results in only one respect: linearity between output and the real exchange rate is rejected. On one hand, the effects of exchange rate changes on output seem to depend on the prevailing monetary regime while on the other hand money affects output differently depending on past exchange rate values.

4. CONCLUSIONS

This paper set out to answer the following questions: First, what is meant by asymmetric behavior of macroeconomic time series? Second, is there any evidence of asymmetric behavior in Swiss macroeconomic data? Third, what might be the possible causes of observed nonlinearities?

The relevance of asymmetry in macroeconomics was motivated by the discussion of alternative models in Section 2, which give rise to asymmetric adjustments in output and prices. As nonlinearity gives rise to various forms of asymmetry, linearity tests can be considered as valuable tests for asymmetry. In order to examine the relevance of asymmetry in Swiss macrodata, I applied a parametric LM test introduced by Luukkonen et al. (1988) to different Swiss macroeconomic time series. To make sure that the tests do not produce spurious results due to certain characteristics of the detrending method used, the results were compared across differently filtered data. I found similar nonlinearities across filters for comparable frequency ranges. Accordingly, linearity is mainly rejected for real investment, industrial production, prices and M1. The evidence of nonlinearity found by the univariate tests indicated the presence of asymmetry in the behavior of Swiss macrodata.

Consequently, I examined the possible source of this phenomenon by extending the univariate LM test in two respects: (1) I used a multivariate version of the LM test as proposed in Granger and Teräsvirta (1993) and (2) I applied tests which allow to distin-
guish between alternative forms of nonlinearity. The multivariate tests suggested that the effects of money on output are nonlinear depending on past values of money, prices and the exchange rate. On one hand the forms of nonlinearity observed, may support the presence of asymmetry due to rigid prices as inherent in menu-cost models. However, not finding any form of nonlinearity in prices with respect to changes in money is at odds with these models. On the other hand, no clear evidence for models of credit constraints has been observed. These results turned out to be robust also, if alternative measures for output and the exchange rate were used. Based on these findings, two questions remain. First, how can we explain nonlinear effects of the lagged explanatory variables on money? May this indicate a nonlinear reaction function of the Central Bank? Second, what is the role of the exchange rate in the nonlinearities on output observed? Surprisingly, nonlinearity in the effects of exchange rate adjustments has not been discussed in the literature so far. However, this source of asymmetry seems to be of major importance for an open economy like Switzerland and appears to yield a rewarding subject to future research.

DATA APPENDIX

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<thead>
<tr>
<th>Variable</th>
<th>Description</th>
<th>Source</th>
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<tbody>
<tr>
<td>Real GDP, (y)</td>
<td>Real gross domestic product (1990=100), seasonally adjusted.</td>
<td>Bundesamt für Statistik</td>
</tr>
<tr>
<td>Investment</td>
<td>Real gross fixed capital formation (1990=100), seasonally adjusted.</td>
<td>Bundesamt für Statistik</td>
</tr>
<tr>
<td>Net exports</td>
<td>Log ratio between real exports and imports in goods and services.</td>
<td>Bundesamt für Statistik</td>
</tr>
<tr>
<td>Industrial production</td>
<td>KOF industry survey: production.</td>
<td>KOF Zürich</td>
</tr>
<tr>
<td>CPI</td>
<td>Consumer price index (1993 Q2=100).</td>
<td>Swiss National Bank</td>
</tr>
<tr>
<td>WPI, (p)</td>
<td>Wholesale price index (1963=100).</td>
<td>Swiss National Bank</td>
</tr>
<tr>
<td>M1, (m)</td>
<td>Definition 1995 without Fürstentum Lichtenstein, seasonally adjusted.</td>
<td>Swiss National Bank</td>
</tr>
<tr>
<td>M2</td>
<td>Definition 1995 without Fürstentum Lichtenstein, seasonally adjusted.</td>
<td>Swiss National Bank</td>
</tr>
<tr>
<td>Credit</td>
<td>Monetary survey - domestic credit: claims on private sector, seasonally</td>
<td>IFS line SWI32D.A</td>
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<tr>
<td>3-Mo Euro rate</td>
<td>Interest rate on 3 month Swiss Franc, deposits in London (end of period).</td>
<td>OECD</td>
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<tr>
<td>Gov. bond yield</td>
<td>Government bond yield source: IFS line SWI61... deposits in London (end</td>
<td>OECD</td>
</tr>
<tr>
<td>Nom. exchange rate, (e)</td>
<td>Effective trade weighted average exchange rate of Swiss Franc.</td>
<td>Swiss National Bank</td>
</tr>
<tr>
<td>Real exchange rate</td>
<td>Effective export weighted exchange rate index (constant prices 1977 Q3=100)</td>
<td>Koff</td>
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Ravn, M.O. and M. Sola (1996), “Asymmetric effects of monetary policy in the U.S.: positive vs. negative or big vs. small?”, Mimeo, University of Aarhus.


SUMMARY

Up until lately, studies on asymmetric business cycles focused on U.S. data and were mainly based on univariate analysis. This paper extends the literature beyond U.S. data in studying Swiss macroeconomic data for which almost no empirical research is available so far. Secondly, this paper applies specific tests to differently detrended data, thus studying the robustness of the results taking the frequency characteristics of the data into account. Thirdly, multivariate tests will be restricted in order to pin down the underlying source of nonlinearity and its transmission channel, focussing mainly on possible nonlinear effects of money supply.

ZUSAMMENFASSUNG

In diesem Papier werden makroökonomische Daten der Schweiz auf Asymmetrien im Konjunkturzyklus untersucht. Die Analyse trägt zur bestehenden Literatur bei, indem die Nichtlinearitätstests auf unterschiedlich trendbereinigte Daten angewendet werden und somit die Robustheit der Ergebnisse überprüft wird. Um die möglichen Ursachen der Nichtlinearität und deren Transmissionskanäle zu finden, werden im multivariaten Test verschiedene Restriktionen untersucht. Hierbei wird in erster Linie auf mögliche nichtlineare Effekte des Geldangebots getestet.

RÉSUMÉ

Cet article examine des données macroéconomiques suisses en vue de détector des asymétries durant le cycle conjoncturel. L’analyse contribue à la littérature existante en appliquant les tests de non-linéarité sur des données libérées de leurs tendances par diverses méthodes, ce qui permet de tester la robustesse des résultats. Différentes restrictions sont examinées dans le test multivariat afin d’identifier les causes de la non-linéarité et leurs canaux de transmission. Une attention particulière est accordée à d’éventuels effets non linéaires de l’offre de monnaie.