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The Effects of the Swiss Currency Floor on the Exchange Rate Pass-Through in Switzerland

A Structural VAR Analysis

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Abstract

The Swiss Franc experienced a substantial appreciation following the financial and the Euro crises in 2008 and 2011. In response to deflationary pressures, the Swiss National Bank set an exchange rate floor of 1.20 Swiss Francs per Euro in September 2011. The floor was discontinued in January 2015. This study investigates the exchange rate pass-through to import and consumer prices in Switzerland, focusing on the period from 2000 to 2019 and on the Swiss Francs per Euro exchange rate. It further investigates the effects of the introduction and discontinuation of the floor on the pass-through. Using monthly macroeconomic data, the analysis is carried out with recursively identified vector autoregressive models. Pass-through effects are quantified through impulse response functions. A baseline analysis ignores the introduction and discontinuation of the floor, a second one analyses their effects on pass-through dynamics by means of an intervention model with time varying parameters. Baseline results show an incomplete but substantial exchange rate pass-through to import prices, a strong pass-through from import to consumer prices, but a weak exchange rate pass-through to consumer prices. Estimates of the intervention model reveal that, while no significant pass-through is found during the floor period, pass-through to import prices is the strongest after the discontinuation of the floor, while that of import prices to consumer prices the weakest. This suggests the floor having been effective in dampening deflation and the pass-through to consumer prices being primarily blocked by a weak pass-through from import to consumer prices after its discontinuation.

Keywords: Exchange rate pass-through, Switzerland, VAR, exchange rate floor, import prices, consumer prices
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1 Introduction

The degree to which exchange rate changes are transmitted to prices of goods and services, generally referred to as the exchange rate pass-through relationship, has been of interest in international Macroeconomics since the breakdown of the Bretton Woods system in 1973 (Stulz, 2007). For monetary policymakers, a solid understanding of pass-through mechanisms is particularly relevant, as the extent to which exchange rate changes are passed to prices impacts both the transmission mechanism of monetary policy and inflation forecasts. For the specific case of Switzerland, this holds even more for mainly two reasons. First, Switzerland is a small open economy with, under normal circumstances, a floating exchange rate. Small open economies often heavily rely on imports but are usually too small to influence global markets. The relative change in import prices caused by an appreciation/depreciation of the domestic currency cannot be counteracted by a change in import prices stemming from an increase/decrease in the small open economy's import demand caused by the appreciation/depreciation, as the demand change is too small relative to global demand. Therefore, the small open economy is expected to carry most of the exchange rate change (An & Wang, 2012). Secondly, the "safe heaven" nature of the Swiss currency makes it vulnerable to periods of increasing appreciation, thereby potentially causing deflationary processes.

The primary mandate of the Swiss National Bank (SNB) is to ensure price stability, while taking due account of economic developments. In so doing, it creates an appropriate environment for economic growth (Federal Act on the Swiss National Bank, 2003). Within its primary mandate, in response to the increasing appreciation of the Swiss Franc caused by both the financial crisis in 2008 and the Euro crisis in 2011, the SNB decided to introduce a minimum exchange rate of 1.20 Swiss Francs per Euro in September 2011. At the time of the introduction, the SNB stated that "the current massive overvaluation of the Swiss Franc poses an acute threat to the Swiss economy and carries the risk of a deflationary development" and committed itself to "enforce this minimum rate with the utmost deteremination." (Swiss National Bank, 2011, p.1). The introduction of the minimum exchange rate represented a transition to an almost fixed exchange rate regime. Approximately five years later, in January 2015, the SNB decided to let the exchange rate float freely again. The timing of the SNB's sudden decision was motivated by changing conditions of international financial markets, which made a strong appreciation of the Swiss Franc less likely and the maintenance of the minimum exchange rate harder to sustain (Jordan, 2016).

The aim of this study is to provide recent empirical evidence on the degree of the exchange rate pass-through (ERPT) to import and consumer prices in Switzerland by focusing on the Swiss Francs per Euro (*EURCHF*) exchange rate and on the introduction and subsequent discontinuation of the *EURCHF* minimum exchange rate.

In a first step, the transmission of *EURCHF* exchange rate and aggregated import price shocks to aggregated import and consumer prices over the period from 2000.01 to 2019.09 (baseline analysis) is examined. The empirical literature includes different methodologies for estimating the pass-through. Campa et al. (2005) or Burstein and Gopinath (2013), for instance, use single-equation models estimated by ordinary least squares (OLS). Using monthly macroeconomic time series data and following Choudhri and Hakura (2015); McCarthy (2007); Stulz (2007) and many others, this study relies on vector autoregressive (VAR) models, which allow to address the endogeneity of the variables[1]. The magnitude of the pass-throughs at different time horizons is then quantified by means of impulse response functions (IRF). A recursive scheme based on a Choleski decomposition is used for identification of the structural shocks. As evidence from VAR models may heavily depend on the model specification, robustness of results is tested both to alternative identification schemes, such as alternative orderings of the variables and generalized IRF, and to an alternative way of dealing with non-stationary variables, namely a vector error correction model (VECM).

In a second step, special attention is given to the two monetary policy interventions carried out by the SNB during the sample period, i.e. the introduction and the subsequent discontinuation of the *EURCHF* minimum exchange rate. Given the nature of the interventions, the hypothesis is that they caused a change in the transmission of *EURCHF* exchange rate shocks to prices, creating three different ERPT regimes. The aim of the study is then to empirically investigate this hypothesis by means of an intervention model, in which the parameters of the VAR model are allowed to change at the dates of the interventions. Based on a structural break analysis of the system implemented through a series of Chow tests, IRF are allowed to change across the three regimes.

For the baseline analysis, the results of the paper suggest that the ERPT to import prices is incomplete but substantial. Moreover, the transmission of import price shocks to consumer prices is found to be surprisingly strong, almost complete. In contrast, exchange rate changes cause only moderate responses in consumer prices. Taken together, the findings suggest the pass-through to consumer prices being mainly blocked by the incomplete ERPT to import prices, that is by sticky import prices. For the intervention model analysis, results seem to confirm the hypothesis of the two interventions having

[1] Exchange rates and prices are believed to be endogenous due to potential macroeconomic shocks simultaneously affecting and determining both.

changed the ERPT dynamics. Compared to baseline results, the ERPT to import prices is weaker in the first regime, i.e. prior to the introduction of the minimum exchange rate, while substantially stronger in the third regime, i.e. after the discontinuation of the minimum exchange rate. The contrary is found for the transmission mechanism of import price shocks to consumer prices, as it appears weaker in the third regime, while equally strong in the first regime. Estimated ERPT to consumer prices is weak for all regimes as in the baseline analysis. The second part of the analysis, therefore, reveals that after the discontinuation of the minimum exchange rate the transmission of exchange rate changes to consumer prices is likely to be primarily blocked by a weak transmission mechanism from import to consumer prices rather than by sticky import prices. One potential explanation proposed for such results is that domestic producers in competition with imported goods decreased the adjustment of their prices in response to the relative price decrease of imported goods following the Swiss Franc appreciation caused by the discontinuation of the minimum exchange rate, thereby decreasing the response in consumer prices[2]. With regards to the minimum exchange rate period, all estimated pass-throughs appear not statistically different from zero. While potentially simply driven by statistical reasons, such as sample size or limited data variation, the insignificance of results could also be interpreted as evidence of the effectiveness of the introduction of the minimum exchange rate in stopping the deflationary process, which was characterising the Swiss economy and driving the substantial pass-through in times proceeding the introduction.

Existing empirical evidence on ERPT is extremely ample. It mainly includes cross country analyses (see, for instance, Choudhri et al., 2005; McCarthy, 2007) using quarterly macroeconomic data (see, amongst others, Choudhri & Hakura, 2015; Hahn, 2003) and implementing OLS (such as in Burstein and Gopinath (2013)) or multivariate time series methods (see, for instance, Cheikh & Louhichi, 2015; Hahn, 2007). However, despite the particular relevance of the pass-through issue for small open economies, empirical evidence for Switzerland is surprisingly scarce. The studies most closely related to this paper are that of Stulz (2007), who estimates a VAR model for Swiss ERPT dynamics over the sample period 1976 to 2004, and of Žídek and Šuterová (2017), who also implement a VAR analysis to estimate the effects of exchange rate shocks on Swiss inflation over the period 2000 to 2016.

This paper contributes to the existing empirical literature in mainly two ways. First, by focusing on the period from 2000 to 2019, it adds to the already existing but scarce empirical evidence on Swiss ERPT dynamics using data covering almost the most recent available time frame. Apart from Žídek and Šuterová (2017)'s work, to the author's

[2] The discontinuation of the minimum *EURCHF* exchange rate resulted in a sharp, unanticipated and permanent appreciation of the Swiss Franc by more than 11% against the Euro (Bonadio et al., 2020, p.1).

knowledge, no other study provides Swiss ERPT estimates on the basis of data covering the last five to ten years. In light of the developments of the Swiss currency over the last ten years, characterised by drastic appreciations, safe heaven capital inflows and the introduction and subsequent discontinuation of a minimum exchange rate versus its major trading currency, the Euro, filling this research gap seems of high interest and relevance. The second contribution of the paper is the analysis of the effects of the two above mentioned monetary policy interventions. Bonadio et al. (2020) exploit the exogenous variation in the *EURCHF* exchange rate caused by the second intervention, the lifting of the cap, to estimate pass-through speed via an event study. Žídek and Šuterová (2017) exploit the interventions of the SNB to investigate the effect on ERPT of switching to a different monetary policy regime by dividing the sample into pre and post exchange rate interventions periods. However, none of these studies looks at both the introduction and the subsequent discontinuation of the minimum *EURCHF* exchange rate separately. To the author's knowledge, no such analysis has been done, at least in the context of ERPT dynamics. Doing so, however, allows to shed light on potential effects of having a capped currency on the transmission mechanism of exchange rate changes to prices. Furthermore, it allows to investigate the effects of interrupting a currency cap. This provides valuable insights also on the effectiveness of such monetary policy actions in providing price stability and containing deflationary (inflationary) processes caused by appreciations (depreciations) of the domestic currency.

The paper is structured as follows. The following section gives an overview of the theoretical and empirical literature on the pass-through issue. Section 3 presents and describes data used. In Section 4, the empirical methodology is set out. Section 5 presents results of the baseline analysis, while Section 6 discusses their robustness. Section 7 presents both the intervention model used to assess potential changes in ERPT mechanisms caused by the two monetary policy interventions and its results. Section 8 summarizes policy implications and the limitations of the study. Section 9 concludes.

2 Theory and empirical evidence

An extensive literature has improved our understanding of the relationship between exchange rates and prices. This includes earlier theoretical work as well as a wide series of empirical studies, ranging from cross country comparisons to single country analyses, from using macro- to microeconomic data and detailed data on good pricing. This section aims to give the necessary overview of the literature in order to correctly interpret and understand the analysis carried out in this study within its relevant context.

2.1 Theoretical background

The fundamental starting point for the economic analysis of the pass-through relationship is the law of one price (LOP). As explained by Goldberg and Knetter (1996), the LOP states that identical products sell for the same common-currency price in different countries. Let $p_{i,D}$ denote the domestic currency price for good i in home country D, while $p_{i,F}$ the foreign currency price for the same good in foreign country F. The LOP holds for good i if:

$$p_{i,D} = p_{i,F} \cdot E \tag{2.1}$$

where E represents the exchange rate of home country D's currency per unit of F's. If the relation given by 2.1 holds, the market for product i is said to be integrated between the two countries. Furthermore, if it holds for all products sold in both countries, then the absolute purchasing power parity (PPP) theory of exchange rates holds between these two countries. Formally:

$$P_D = P_F \cdot E \tag{2.2}$$

where P_D and P_F are price levels in countries D and F respectively (pp.1245-1246). Assuming perfect competition in both countries, price levels equal marginal costs in the respective currencies of the countries. Under this condition, if the exchange rate changes and the foreign price level remains unchanged, the domestic price level changes one to one. The ERPT is then said to be complete (Mann, 1986, p.367).

Ample empirical evidence has suggested that the ERPT is rarely complete, indicating that local currency prices of foreign products tend not to fully respond to exchange rate changes (Fleer et al., 2016, p.2). In light of this, theoretical research on the relationship between exchange rates and prices has mainly been focused on explaining different degrees of pass-through incompleteness. An early branch of literature has tried to accomplish this by looking at the pricing behaviour of firms. Another, more recent, branch of literature has tried to complement the former by considering macroeconomic forces as well.

2.1.1 Pricing behaviour of firms

The earliest theoretical literature investigating pass-through relationships adopted a microeconomic approach and examined the incompleteness of the pass-through in the context of imperfect competition, usually introduced in form of market segmentation or product differentiation, which enable exporting firms to differentiate prices across countries. Following Mann (1986), in the context of imperfect competition, the concept of pass-through is given by:

$$\Delta P_D = \Delta P_F + \Delta E = \Delta C_F + \Delta M_F + \Delta E \tag{2.3}$$

where C_f and M_f are costs and margin over costs in foreign currency. Then, according to 2.3, deviations from a complete pass-through can arise because of differences in the cost of supplying the good to different locations or because firms discriminate prices across locations by charging different markups (Burstein & Gopinath, 2013, p.391). Indeed, Krugman (1986) introduced the concept of pricing-to-market (PTM), according to which exporting firms adjust their destination-specific markups in order to compensate for exchange rate changes and allow for an optimal price adjustment. Goldberg and Knetter (1996) further develop the concept of PTM and confirm Krugman (1986)'s theory by showing that, provided firms have some market power to discriminate prices across destination markets, PTM represents a valid explanation for incomplete ERPT to import prices.

From a macroeconomic perspective, the study of pass-through incompleteness is a more novel field of research. Traditional macroeconomic assumptions of perfect price competition and fully flexible prices in both domestic and foreign markets imply absolute PPP and thus complete pass-through. Obstfeld and Rogoff (1995) are the first to include price rigidities and market imperfections in their dynamic general equilibrium model. However, in their model, nominal prices are set in the currency of the producing country. Therefore, despite the presence of one period price rigidities, nominal exchange rate fluctuations still cause one to one reactions in prices of imported goods, implying a complete pass-through. An incomplete pass-through was only later allowed for by Betts and Devereux (2000), who extended the model of Obstfeld and Rogoff (1995) by introducing PTM. More precisely, they modified the pricing behaviour of firms such that some firms set their nominal prices in the currency of the destination country instead of in their producing country's one. Combined with the presence of one period nominal price rigidities, this extension results in an incomplete pass-through in the short run, as changes in the exchange rate cannot be immediately transmitted to prices set in the destination country's currency. The degree of incompleteness then depends on the share of firms that set their prices in the destination country's currency: the bigger this share, the lower the pass-through to domestic prices in the destination country.

As noted by Stulz (2007), these two early models ignore some relevant aspects. Most importantly, none of the two distinguishes between different stages of the distribution chain. If imports partly consist of intermediate goods that undergo non-traded production or distribution processes before being consumed, the ERPT may be dampened by such production or distribution channels (p.7). The pass-through might then be incomplete even in the case of producing country currency pricing (see, for instance, the model provided by McCallum and Nelson (1999)).

As pointed out by Mann (1986), the pricing behaviour of firms and market structures interact with and are affected by macroeconomic forces and uncertainty. Therefore, attention has also been dedicated to which macroeconomic factors affect the pass-through relationship the most.

2.1.2 Macroeconomic factors

Given the relevance of macroeconomic forces for the behaviour of firms and price development, the literature has identified following key macroeconomic factors related to the ERPT: the size of a country, its trade openness, the exchange rate volatility and persistence, the inflation and monetary policy environment.

An and Wang (2012) explain how, in theory, the size of a country, measured by real GDP, is inversely related to the completeness of the pass-through for two main reasons. First, the deflationary effect of an appreciation is likely to increase the import demand of the country. If the economy is large enough to influence the global market, world's price of imports will rise, increasing the economy's price level and thus counteracting the price decrease caused by the appreciation, thereby decreasing the measured pass-through. Second, when it comes to big markets, foreign exporters are more willing to adjust markups rather than prices, since in such markets they have greater incentives to maintain their market shares. With regards to Switzerland, however, it seems reasonable to assume that its size and market share are too small to impact the world's import demand and supply as well as foreign exporters's incentives. Consequently, the Swiss economy is expected to carry most of the exchange rate changes, thereby increasing the measured pass-through.

Trade openness of a country, usually measured as the import penetration ratio[1], is expected to be positively correlated with the degree of pass-through: a large import penetration may imply less competition from domestic producers, allowing foreign companies

[1] The import penetration ratio represents the participation of foreign firms in the domestic economy and is measured by the share of imports in domestic consumption.

to pass exchange rate changes to prices of importing countries rather than forgoing on profit margins (Cheikh & Rault, 2016, p.84).

Exchange rate volatility and the persistence of exchange rate shocks as important macroeconomic determinants of the pass-through degree have been widely analysed in the literature (see, for instance, Corsetti et al., 2008; Mann, 1986; Žídek & Šuterová, 2017). With regards to exchange rate volatility, Mann (1986) argues how, based on the PTM principle, exchange rate volatility is inversely related to the degree of pass-through. Greater exchange rate volatility may make importers reluctant to continuously pass changes in the exchange rate to consumers in destination markets by adjusting prices. Instead, they are much more incentivised to continuously adjust profit margins, thereby reducing measured pass-through. The persistence of exchange rate movements is supposed to affect the intensity of the pass-through in a similar way as the exchange rate volatility. If firms expect a change in the exchange rate to last for a long period, they are more likely to pass the shock to prices and consumers (An & Wang, 2012, p.6). In line with this, Jašová et al. (2019) explain how larger exchange rate movements have a higher chance to overcome menu costs of price changes and, thus, are more likely to be passed-through to consumer prices (p.2).

A great amount of theoretical and empirical studies on the link between inflation environment and the ERPT have provided strong evidence for the pass-through being weaker in environments with low and stable inflation and inflation expectations. The Bank of Canada (2000) was among the first to connect a weak pass-through to a low inflation environment. In one of its Monetary Policy Reports in 2000 it claims "The low-inflation environment itself is changing price-setting behaviour. When inflation is low and the central bank's commitment to keeping it low is highly credible, firms are less inclined to quickly pass higher costs on to consumers in the form of higher prices." (p.9). Afterwards, Taylor (2000) developed a theoretical framework according to which the degree of pass-through depends on the inflation environment and provided empirical evidence based on the US economy supporting his theory. Drawing on Taylor (2000)'s work, the link between pass-through and inflation has been empirically examined in many other studies. To mention one, Gagnon and Ihrig (2004) implemented a cross-country analysis comprising twenty industrialised countries and found that the ERPT declined in most countries where, in the early 1990s, there had been a regime shift towards more inflation stabilisation. In their model, when the monetary authority focuses strongly on stabilising inflation, the pass-through of exchange rate movements to consumer prices is lower (p.316). Stulz (2007) found similar results for Switzerland.

2.2 Existing empirical evidence

Besides theoretical work, there exists a wide and established empirical literature on the topic of ERPT. It encompasses different types of studies, ranging from cross country analyses to single country ones, implemented with macroeconomic or detailed micro-level data. With regards to methodologies used, the studies vary from multivariate time series methods (mainly VAR models) to univariate ones such as linear OLS or event study approaches.

Menon (1995) was one of the first to provide a literature review on the topic. Looking at 43 studies, he finds an incomplete pass-through to be a common feature for the majority of the countries studied. Furthermore, the pass-through degree seems to differ across countries and, for the same country, across different studies. He traces these differences back to differences in empirical methodologies, data and variables included. In line with the price discriminating behaviour of firms, he also finds variation in the degree of pass-through across products and markets.

Through his cross country study, McCarthy (2007) was one of the first to introduce VAR analysis to estimate the pass-through mechanism along the distribution chain. He analyses nine developed countries, including Switzerland. In general, he finds a declining pass-through along the distribution chain, with a modest ERPT to consumer prices and a more substantial one to import prices. In line with theory, he finds import share and exchange rate persistence of a country to be positively correlated with the intensity of the pass-through, while exchange rate volatility to be negatively correlated. For Switzerland, he reports a substantial ERPT to import prices but lower relative to the other industrialized countries. With regards to Swiss consumer prices, he finds almost no pass-through.

That the pass-through declines along the distribution chain has been found in many other similar studies. Hahn (2003), for instance, finds a much higher pass-through to import prices than to consumer prices when estimating ERPT in the Euro area using a VAR model and quarterly data. Choudhri et al. (2005) find similar results when looking at six G7 countries (namely Canada, France, Germany, Italy, Japan and UK). By means of predicted responses based on a quantitative model as well as VAR analysis, they find average ERPT to import prices of 0.45, 0.73 and 0.22 percent after 1, 4, and 10 quarters respectively following a one percent change in the exchange rate while, after the same number of quarters, of only 0.02, 0.11 and 0.19 to consumer prices.

Choudhri and Hakura (2015) estimate pass-throughs for several countries based on OLS as well as VAR models. Similarly to the other mentioned studies, using data on 18 advanced economies and 16 emerging ones, they find an incomplete but substantial ERPT to import prices. Average pass-throughs for advanced economies to import prices amount to 0.67 and 0.60 percent (following a one percent change in the exchange rate) when using OLS and VAR models respectively. For Switzerland, they estimate an OLS pass-through to import prices of 0.63 and of 0.52 when using a VAR model, both one quarter after the exchange rate shock.

Also by means of a multivariate time series model, Cheikh and Louhichi (2015) find a declining ERPT along the distribution chain when analysis 12 Euro area countries. Despite some differences across countries, the lowest effect is found on consumer prices for all countries. Furthermore, when trying to understand drivers of country differences, they find inflation level, inflation volatility and exchange rate persistence to be the main macroeconomic factors influencing the pass-through, thereby confirming theoretical considerations.

When univariate linear models are used instead, the standard approach for estimating ERPT has been to regress changes in some measure of domestic prices on past and present changes in the exchange rate and additional control variables. An example of such studies is the one of Campa and Goldberg (2005), who look at 23 different countries and estimate an average short-run (one month) pass-through to import prices of 0.46 percentage points after a one percent change in the exchange rate and a long run (four months) one of 0.64. For Switzerland specifically, they find a short run pass-through to import prices of 0.68 and an almost complete one in the long run of 0.93. However, as pointed out by Forbes et al. (2018), such linear models seem to poorly address the simultaneity of exchange rates and prices.

A more recent strand of literature has started to use extensive micro-level data on the pricing decisions of firms to investigate factors found in the microeconomic theory of ERPT, such as price adjustment frequency (see Gopinath & Itskhoki, 2010), the role of currency of pricing (see, for instance, Devereux et al., 2015), the role of mark-up adjustments and local costs (Nakamura & Zerom, 2010), the distribution of price changes (Berger & Vavra, 2013) and the intensity of competition in final product markets (Amiti et al., 2016).

In essence, what emerges from the empirical literature is that ERPT seems to be incomplete. With that said, there is a lot of variation across countries, with the ERPT ranging from being substantial to quite low depending on both macro- and microeconomic factors. Among these factors, inflation environment and exchange rate volatility seem to be the most relevant, thereby confirming theoretical considerations. However, they are also the most analysed by the literature. Finally, ERPT seems to gradually weaken along the distribution chain.

Apart from cross country studies, with regards to Switzerland specifically, despite the particular relevance of pass-through issues for small open economies, the empirical evidence is surprisingly sparse. Amongst the few studies, the one of Herger (2012) endeavours to estimate pass-through elasticities of Swiss import prices as linear OLS coefficients using monthly macroeconomic data covering the 1999 to 2010 period. His results suggest a highly incomplete pass-through of around 0.30 percent on aggregate, which, however, varies a lot across industries. More recent studies looked at microeconomic factors such as currency of invoicing and price dispersion (see Bonadio et al., 2020; Fleer et al., 2016; Sarah et al., 2017). Using linear univariate regression models and detailed micro-level data, they all find a significant but incomplete pass-through to Swiss import prices. Finally, this study is most closely related to the ones of Stulz (2007) and Žídek and Šuterová (2017). By means of a VAR model for Switzerland, Stulz (2007) estimates the ERPT along the distribution chain using monthly data from 1976 to 2004. He finds an import price adjustment to a one percent shock in the exchange rate of about 0.50 percent after one year and a consumer price adjustment of only 0.18 percent after two years. Despite the similarity of Stulz (2007)'s analysis to this one, results are expected to differ for mainly two reasons. First, contrarily to the analysis of this study, rather than focusing on the *EURCHF* exchange rate, Stulz (2007) looks at changes in the effective general nominal exchange rate[2]. Second, he looks at an older sample period in which no drastic monetary policy intervention in the foreign exchange rate market had happened yet.

Žídek and Šuterová (2017), on the contrary, look at the more recent period between 2000 and 2016. They, too, estimate ERPT along the distribution chain by means of a VAR model for Switzerland. Mostly connected to this study, however, is their aim to evaluate the effect of exchange rate interventions of the SNB on consumer prices by dividing the sample into pre and post interventions periods. Their results suggest that the exchange rate interventions did enhance the ERPT into inflation (i.e. consumer prices). Other than using quarterly instead of monthly data, Žídek and Šuterová (2017)'s study mainly differs from this one in that it looks at the effect of the two interventions combined. That is, they treat the two interventions as a single switch into a new monetary policy regime and try to analyse its effect on the ERPT. This study, in contrast, aims to look at the effect of the two interventions separately in order to assess the impact of introducing and subsequently lifting a cap on the currency.

[2] The effective nominal exchange rate used by Stulz (2007) is computed by considering the 24 biggest trading partners of Switzerland.

3 Data

Many empirical studies have analysed ERPT elasticities by simulating the dynamics of small open economies through VAR models. This study aims to investigate the effects of fluctuations in the *EURCHF* nominal exchange rate on Swiss import and consumer prices by estimating a VAR model for the small open economy of Switzerland. In this section, the choice of the variables to be included in the model and their statistical characteristics are described.

3.1 Data description and transformation

Data used for this study includes monthly macroeconomic time series covering the period from 2000:01 to 2019:09. When estimating ERPT via impulse responses of a VAR model, one would like to include those variables of the small open economy under investigation that are the most relevant for ERPT dynamics. To this end, this study closely follows the information set included by Choudhri et al. (2005), who estimate ERPTs via VAR models for six open economies, and by Stulz (2007), who estimates a similar model for Switzerland.

To capture real economic activity, demand shocks and demand fluctuations over the business cycle, a measure for the Swiss output gap is included (gdp_t). The purpose of this variable is to control for the state of the economy, i.e. whether it is in a recession or in an expansion period[1]. Controlling for the state of the economy is of relevance, as a recession is assumed to be associated with a lower ERPT. If demand is weak, firms might be more afraid to increase prices in response to a currency depreciation due to higher risk of losing market share (Mann, 1986). Under the assumption that real output fluctuates around some potential level, the monthly output gap can be defined as the log deviation from the actual monthly real GDP to its potential level. In this study, gdp_t is then computed by means of the Hodrick-Prescott filter method, which considers deviations of actual GDP from potential GDP as deviations from its trend component (computed by the filter) (Stulz, 2007, p.9).

Given the role monetary policy plays on the price level of the economy, a variable capturing the effects of monetary policy is to be included in the system. Some studies have included short-term interest rates (see, for instance An & Wang, 2012; Hahn, 2003; McCarthy, 2007), others a measure of a monetary aggregate (as done by Stulz (2007)). In

[1] A negative (positive) output gap is assumed to imply that the economy is in a recession (expansion).

light of the interest rate parity (IRP)[2] theory and the fact that policy rates represent one of the main monetary policy instruments of central banks, interest rates might seem to be the most appropriate variable to control for monetary policy actions. However, conventional theory and monetary policy tools might not apply as well as usual to the period considered here and to the Swiss Franc. Indeed, citing the Chairman of the Governing Board of the SNB: "In normal times, if a country's currency appreciates to the point where it begins to jeopardise price stability, that country's central bank will react by lowering the reference interest rate. Lowering rates makes portfolio investments denominated in the relevant currency less attractive, thereby weakening it on the foreign exchange markets. However, when nominal interest rates reach their lower bound at or close to zero, central banks must take unconventional measures." (Jordan, 2016, p.4). The recession caused by the financial crisis of 2008 led the SNB to decrease its policy rate to zero. At the same time, despite interest rates being at the zero bound, international economic uncertainty combined with the Swiss Franc being seen as a "safe heaven" currency resulted in a substantial appreciation of the Swiss Franc against the Euro. Therefore, to influence the exchange rate and to introduce and maintain the floor of 1.20 Swiss Francs per Euro, the SNB had to recur to the unconventional tool of using the monetary base. More precisely, it started increasing the Swiss monetary base in order to buy enough Euros to maintain the floor. In fact, increasing liquidity was one of the main reasons for which the floor was later lifted. Again citing Jordan (2016): "Given the changed international environment, the minimum exchange rate of 1.20 Francs per Euro was no longer sustainable. The SNB could only have enforced it through ongoing foreign currency purchases of rapidly increasing magnitude – in an environment where there was no prospect of a long-term stabilisation on the exchange rate front. Delaying the decision to discontinue would not have mitigated the economic consequences, but the losses for the SNB would have been far greater." (p.6). In essence: monetary policy actions aimed at affecting the Swiss currency during the period considered were implemented mostly through changes in the monetary base. For this reason, in its main specification, this study uses a measure of a monetary aggregate to control for the effects of monetary policy. More precisely, following Stulz (2007), the monetary aggregate M3 is used, which includes currency in circulation, sight deposits, time deposits and savings deposits (Swiss National Bank, 2007, p.9). However, sensitivity of results to this choice is analysed by estimating the model also with various interest rates as the monetary policy variable. As shown in Section 6, results are little affected by changing the type of monetary policy variable.

[2] Similarly as the LOP and the PPP, the IRP theory implies a no arbitrage condition where investors are indifferent to interest rates available in bank deposits of two countries as the expected return on domestic assets will equal the exchange rate adjusted expected return on the foreign assets (Engel, 2013).

At the centre of the empirical analysis lies the nominal exchange rate of the Swiss Franc versus the Euro (e_t) (where an increase denotes a depreciation of the Swiss Franc), together with the Swiss import price index (IPI) (ipi_t) and the Swiss consumer price index (CPI) (cpi_t).

Finally, in order to control for changes in consumer prices abroad, which might be independent from fluctuations in the exchange rate but still affect import and consumer prices in Switzerland, a variable of foreign consumer prices must be included. As the analysis focuses solely on the transmission of changes in the *EURCHF* exchange rate on Swiss prices, the harmonized consumer price index (HCPI) ($hcpi_t$) of the Euro area is chosen.

Time series used were retrieved from the SNB database, the Swiss Federal Statistical Office (SFSO), the Swiss State Secretariat for Economic Affairs (SSEA), the Swiss Federal Custom Administration (SFCA) and the European Central Bank (ECB) database. The sample period covers monthly observations from 2000:01 to 2019:09, thus including a total of 237 observations and representing different exchange rate regimes: until October 2011 the *EURCHF* exchange rate was fully flexible, from September 2011 to December 2014 it was artificially maintained at a minimum of 1.20 and, finally, from January 2015 until the end of the sample it was again free to float. Monthly data rather than quarterly data was chosen as the main interest of the study lies in short-run pass-through dynamics. Given the importance of reliable short-term inflation forecasts for monetary policy makers, short-term dynamics seem to be the most policy relevant.

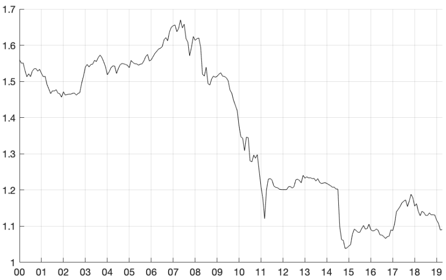
All series are transformed in their natural logarithm (except for the output gap which, by definition, is already in logarithmic form). Furthermore, all series except the nominal exchange rate are seasonally adjusted by means of the Census-X-12 procedure[3].

Figure 3.1 depicts the exchange rate, the IPI and CPI in levels over the sample period. From a visual inspection of the *EURCHF* exchange rate series (panel a), the substantial and continuous appreciation after the financial crisis of 2008 is quite visible. So is a sudden interruption of the appreciation around 2011 (i.e. with the introduction of the minimum exchange rate) and another appreciation shock around 2015 (i.e. with the interruption of the minimum exchange rate). The interventions are not as clearly visible in the IPI and CPI graphs. However, both the IPI and the CPI seem to show a drop around 2015, which might coincide with the interruption of the minimum exchange rate.

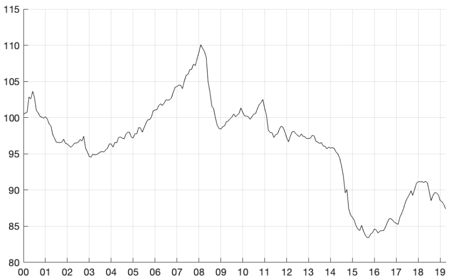
[3] The Census-X-12 method (X-12-ARIMA method) was developed by the United States (US) Bureau of the Census in 1998, which made the X-13ARIMA-SEATS program publicly available for its implementation. In essence, the program fits a series of autoregressive integrated moving average (ARIMA) processes on the time series to decompose it into three components: systematic calendar-related effects, irregular fluctuations, and trend behaviour. The first two together make the seasonal fluctuations, the trend component is the one of interest kept for the analysis (U.S. Census Bureau, 2002). In this study, the program has been implemented through the Matlab toolbox *X-13 Toolbox for Seasonal Filtering* developed by Yvan Lengwiler (2020).

A detailed structural break analysis of these three series is carried out in Section 7.1. Summary statistics for each series are depicted in Table A.1 in Appendix A.1, together with graphical representations of all series.

Figure 3.1: Exchange Rate, Import Price Index and Consumer Price Index



(a) *EURCHF* Nominal Exchange Rate



(b) Import Price Index

(c) Consumer Price Index

Notes: The figure depicts monthly time series in levels for the *EURCHF* nominal exchange rate (panel a), the Swiss Import Price Index (panel b) and the Swiss Consumer Price Index (panel c) from 2000:01 to 2019:09. Import Price Index and Consumer Price Index are seasonally adjusted by means of the Census-X-12 procedure. Sources: Author's rendering of SNB and SFSO data (2020).

Of the time series used, Swiss GDP is the only one that does not exist at monthly frequency but is only available in quarterly observations. Therefore, monthly values have to be estimated from the available quarterly ones. Based on Chow and Lin (1971)'s method to disaggregate quarterly data into monthly frequency, this is done by using GDP related series available at monthly frequency, of which a linear combination is assumed to be highly correlated with the true monthly GDP observations. As suggested by Cuche and Hess (2000) and implemented by Stulz (2007), the choice of the related series can be based on the components of the expenditure side of the GDP[4]. However, as there are only

[4] The expenditure side of the GDP is given by private consumption, private domestic investment, government expenses and net exports (Cuche & Hess, 2000, p.168).

a few series with monthly frequency available for Switzerland, one needs to find proxies for some of the components (Cuche & Hess, 2000, p.13). I, therefore, use monthly data on retail sales to proxy for private consumption, monthly data on imports of investment goods as a proxy for total imports and investment and, finally, monthly data on exports of all goods as a proxy for total exports. Government expenditures are excluded as not only no sensible proxy at monthly level is available for them, but also because of their low covariance with the business cycle due to tied up spendings (Cuche & Hess, 2000, p.15). The chosen related series were collected for the sample period 2000:01 to 2019:09, seasonally adjusted by means of the Census-X-12 procedure and used as high frequency indicators to disaggregate the quarterly GDP data from 2000Q1 to 2019Q3, as provided by the SSEA. More details on this procedure as well as on the validity of its results are discussed in Appendix A.2.

3.2 Pretesting of variables

When working with time series one would like to investigate on their statistical characteristics, especially with regards to stationarity and the presence of unit roots. This is of relevance when evaluating if and how the series are to be transformed before being included in the VAR model.

To test for unit roots, an Augmented Dickey-Fuller (ADF) as well as a Phillip-Perron (PP) test have been implemented on all series. Following Enders (2010)'s approach, each series has been modelled as an autoregressive process. More precisely, depending on the mean of the series and whether or not it depicts a trend, the autoregressive process is given by one of the following three (pp.206-207):

$$\Delta y_t = \gamma y_{t-1} + \sum_{i=1}^{k} \Delta y_{t-i+1} + \epsilon_t \qquad \text{(No drift, no trend)}$$

$$\Delta y_t = a_0 + \gamma y_{t-1} + \sum_{i=1}^{k} \Delta y_{t-i+1} + \epsilon_t \qquad \text{(Drift, no trend)}$$

$$\Delta y_t = a_0 + \gamma y_{t-1} + a_2 t + \sum_{i=1}^{k} \Delta y_{t-i+1} + \epsilon_t \qquad \text{(Trend)}$$

where y_t represents the series tested. Visual inspection of the data helped in choosing the type of regression to be used for the tests. Optimal number of lags to be included in each regression is selected by means of the Akaike Information Criterion (AIC) and the Bayesian Information Criterion (BIC). These have been computed for AR(p) models fitted on each series, with p ranging from zero to $p_{max} = 15^5$ (whenever the two criterias differ, I chose the number of lags selected by the BIC, as it is considered to be more parsimonious (Lütkepohl, 2005)). Before estimating test regressions with $k = p - 1$ augmentations, a final Ljung-Box test is computed on the residuals of the selected AR(p) models to check if they are serially uncorrelated. If the null hypothesis of zero correlation in the residuals cannot be rejected, the ADF and PP tests are computed, otherwise lags are added to the $AR(p)$ model until the test fails to reject it. As a further test to investigate the presence of unit roots, besides the ADF and PP tests, a Kwiatkowski, Phillips, Schmidt and Shin (KPSS) test is also computed.

Table 3.1 depicts, for each time series, test results as well as regression model chosen for the implementation of the tests and lags selected. Note how under the null hypotheses of the ADF and PP tests, the series contains a unit root, whereas under the null hypothesis of the KPSS test, the series does not contain a unit root (Franke et al., 2015, p.248).

Results shown in Table 3.1 are in line with expectations: for all series except for the output gap, the hypothesis of the presence of a unit root cannot be rejected, i.e. all series except for the output gap seem to be nonstationary. For the output gap, on the contrary, the null hypothesis of the presence of a unit root implied by the ADF and by the PP tests can be rejected at the 1% significance level. Similarly, KPSS test results indicate that the null hypothesis of stationarity cannot be rejected. This too is very much in line with expectations, as the output gap, per definition, should be stationary.

Table 3.1: Unit Root Testing Results

Variable	**Transformation**	**Model**	**Lags** p	**ADF Test p-value**	**PP Test p-value**	**KPSS Test p-value**
Exchange Rate (e_t)	log-level	Drift, no trend	2	0.861	0.881	0.01
IPI (ipi_t)	log-level	Drift, no trend	3	0.696	0.879	0.01
CPI (cpi_t)	log-level	Trend	1	0.886	0.886	0.01
Output Gap (gdp_t)	level	No drift, no trend	5	0.001	0.001	0.1
M3 (m_t)	log-level	Trend	2	0.930	0.923	0.01
HCPI ($hcpi_t$)	log-level	Trend	15	0.889	0.780	0

Notes: The table depicts results of unit root testing for each series of Augmented Dickey-Fuller (ADF) , Phillip-Perron (PP) and and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) tests. Lags are selected from estimating AR(p) models with p ranging from 0 to 15 on the basis of the AIC and BIC. Whenever the two criteria differed, lags selected by the BIC were chosen.

[5] p_{max} has been defined according to the following "rule of thumb": $p_{max} = q \cdot \left(\frac{T}{100}\right)^{\frac{1}{4}}$, where T is the sample size and q the number of times in which a year is divided, i.e. 12 when using monthly data (Hashimzade, 2013, p.69).

4 Empirical methodology

The aim of this study is to estimate the *EURCHF* ERPT by means of impulse response functions of a VAR model for Switzerland. The choice of the model is in line with the standard empirical literature on ERPT estimation (see, for instance An & Wang, 2012; Cheikh & Louhichi, 2015; Hahn, 2003; McCarthy, 2007; Stulz, 2007). Indeed, VAR models are well suited for such estimation for a number of reasons. First, compared to univariate models, VAR models allow for a greater interaction across the variables included. Given the transmission mechanisms to be estimated and the underlying macroeconomic relationships, greater interaction is a desirable feature. Second, as mentioned by An and Wang (2012), VAR models better address the endogeneity problem inherent in single-equation regressions (pp.2-3). Furthermore, as highlighted in the literature review, VAR models allow to estimate the ERPT to different prices along the distribution chain, such as import prices and consumer prices (see, amongst others, An & Wang, 2012; Choudhri et al., 2005; Hahn, 2003; Stulz, 2007).

The ability to estimate the pass-through along the distribution chain is of most importance for a variety of reasons. As explained in Section 2.1.2, import penetration is assumed to be positively correlated with the pass-through intensity. In line with this, exchange rate movements are transmitted to consumer prices through two channels: first, through changes in the prices of imported goods, and, second, through changes in the prices of domestically produced goods in response to price changes of imported goods. The extent of the pass-through to consumer goods will therefore depend on the ERPT to import prices, the share of imports in the basket of consumer goods, and the reaction of prices of domestically produced goods to exchange rate movements (i.e. to changes in prices of imported goods) (An & Wang, 2012, p.7). Estimating the pass-through along the distribution chain therefore provides insights on each of the two channels. Moreover, as previously hint at, it seems reasonable to assume that not all imported goods enter the domestic economy as final consumption goods. Some imported goods are likely to be intermediate goods that need to undergo further production processes (raw materials would be an example of such) or distribution processes before being consumed. As costs of such processes would be in the domestic currency, they would impact the final consumption price independently of an exchange rate change, thereby dampening the pass-through to final consumer goods' prices (Stulz, 2007, p.7).

As a first step, this section derives the baseline VAR model for the Swiss economy. In a second step, the identification strategy is discussed. Finally, the derivation of the IRF and the estimation method are explained, as well as the choice of the model's lag order.

4.1 The vector autoregressive model

VAR models are used for multivariate time series, where each variable is a linear function of past lags of itself and past lags of the other variables. Following Lütkepohl (2005), the reduced form of the VARX(p, q)[1] model of interest may be written as:

$$y_t = \nu + A_i y_{t-1} + A_2 y_{t-2} + ... + A_p y_{t-p} + B_0 x_t + B_1 x_{t-1} + ... + B_q x_{t-q} + u_t \tag{4.1}$$

where $y_t = (y_{1t}, ..., y_{Kt})'$ is a $(K \times 1)$ vector containing the values of K endogenous variables at time t, $x_t = (x_{1t}, ..., x_{Mt})'$ is a $(M \times 1)$ vector containing the values of M exogenous variables [2] at time t, A_i and B_i are fixed $(K \times K)$ and $(K \times M)$ coefficient matrices and $\nu = (\nu_1, ..., \nu_K)'$ is a fixed $(K \times 1)$ vector of intercept terms allowing for the possibility of a nonzero mean $E(y_t)$. Finally, $u_t = (u_{1t}, ..., u_{Kt})'$ is a $(K \times 1)$ standard white noise innovation process. That is, $E(u_t) = 0, E(u_t u_t') = \Sigma_u$ and $E(u_t u_s') = 0$ for all $s \neq t$. Therefore, Σ_u is not diagonal and error terms u_t are correlated with each other.

Here, t corresponds to a month and the endogenous variables included are the following five (in natural logarithms): the output gap (gdp_t), monetary aggregate M3 (m_t), the $EURCHF$ nominal exchange rate (e_t), the Swiss IPI (ipi_t) and the Swiss CPI (cpi_t). To control for the effect of price movements abroad, the consumer price level of the Euro area, proxied by the HCPI ($hcpi_t$), is included as exogenous variable. The foreign price level is modelled as exogenous as it is assumed that a small open economy like Switzerland has no influence on the foreign price level (Stulz, 2007, p.9).

As revealed in Section 3.2, pretesting procedures provided evidence for all series except the output gap to be integrated of order one. According to Enders (2010), there is an issue of whether variables included in a VAR model need to be stationary or not. Sims et al. (1990), for instance, recommend against differencing even if the variables contain a unit root, as the purpose of a VAR analysis is to investigate on the interrelationships among the variables. Differencing could then misleadingly "throw-away" information regarding potential comovements in the data, in which case the use of a vector error correction model (VECM) instead might be preferable. However, if the variables are not cointegrated (see Section 6.3 for an explanation of cointegration), it is more advisable to estimate a VAR in first differences (Enders, 2010, pp.291,384). As the interest of this analysis lies in short-term dynamics, and in light of the fact that many other similar studies have estimated VAR models in first differences (see, amongst others, Choudhri et

[1] VARX refers to a VAR model that includes also exogenous variables with q lags.

[2] Here, exogenous refers to variables that are not determined within the system and on which one can condition the analysis (Lütkepohl, 2005, p.388).

al., 2005; Hahn, 2003; McCarthy, 2007; Stulz, 2007), to address potential nonstationarity in the data, all variables integrated of order one are included in first differences. The alternative strategy for dealing with nonstationarity of testing for cointegration and using a VECM is implemented as a robustness check in Section 6.3.

As all variables are in natural logarithms, inserting the nominal exchange rate, the IPI and the CPI in first differences leads to an economically meaningful interpretation of the IRF, namely of pass-through elasticities. More precisely, impulse responses represent the percentage change in the relevant price index following a one percentage point shock in the nominal exchange rate or in import prices.

Formally, y_t and x_t are the following:

$$y_t = (gdp_t, \Delta m_t, \Delta e_t, \Delta ipi_t, \Delta cpi_t)'$$
$$x_t = (\Delta hcpi_t)$$

where Δ denotes first differences. Following Lütkepohl (2005), to simplify notation, the VARX(p, q) model can be written in its lag operator form (p.22):

$$A(L)y_t = \nu + B(L)x_t + u_t$$

where $A(L) = (I_K - A_1L - ... - A_pL^p)$ and $B(L) = (B_0 + B_1L + ... + B_qL^q)$.

In this framework, impulse response analysis can be seen as a counterfactual experiment of tracing the marginal effect of a shock to one variable through the system by setting one component of the error term u_t to one and all other components to zero and evaluating the responses of the endogenous variables at time t to such an impulse as time goes by. However, because the components of u_t are instantaneously correlated (as mentioned above, Σ_u is not diagonal), such a counterfactual experiment may not properly reflect the actual responses of the economic system of interest. Due to the instantaneous correlation, an impulse in one variable is likely to be accompanied by an impulse in another one and, therefore, cannot be considered in isolation (Lütkepohl, 2018, p.2). For this reason, prediction errors u_t need to be transformed into economically meaningful innovations by computing orthogonalized impulse responses. This can be done by deriving the structural representation of the model with corresponding structural innovation terms ε_t that are instantaneously uncorrelated. As the identification of the structural form and its innovations are not affected by deterministic terms, to further simplify notation, these will be ignored in the subsequent derivations.

If $A(L)$ is invertible, there exists a lag operator $\Phi(L)$ such that $\Phi(L) = A(L)^{-1}$. It then follows not only that the VAR(p) process is stationary, but also that there exists a moving average (MA) representation of it (Lütkepohl, 2005, p.25). Thus, assuming invertibility of $A(L)$, the MA representation of 4.1 can be derived by premultiplying 4.1 by $\Phi(L)$:

$$y_t = \Phi(L)u_t = \sum_{i=0}^{\infty} \Phi_i u_{t-i} \tag{4.2}$$

where $\Phi_0 = I_k$ and Φ_i represent the MA coefficient matrices, which contain impulse responses from shocks in the error terms u_t.

The structural representation of 4.1 can now be constructed by premultiplying the llag operator form of 4.1 by the matrix B so that

$$A^*(L)y_t = \varepsilon_t \tag{4.3}$$

where $A^*(L) = B \cdot A(L)$ and $\varepsilon_t = B \cdot u_t$. The economically significant innovations ε_t are assumed to be standard white noise with variance of one, so that their variance covariance matrix can be normalized to I_K. Formally: $\Sigma_\varepsilon = E(\varepsilon_t \varepsilon_t') = I_K$. As before, assuming invertibility of $A^*(L)$, the MA representation of 4.3 can be derived by premultiplying by $\Theta(L) = A^*(L)^{-1}$:

$$y_t = \Theta(L)\varepsilon_t = \sum_{i=0}^{\infty} \Theta_i \varepsilon_t \tag{4.4}$$

where Θ_i now contains the impulse responses from shocks in ε_t. As $\Sigma_\varepsilon = I_K$, the components of ε_t are instantaneously uncorrelated and Θ_i therefore contains orthogonalized impulse responses.

As innovations ε_t are not observed, they need to be identified. Identification can be achieved by estimating error terms u_t and variance covariance matrix Σ_u by means of OLS from 4.1 (for details on the estimation procedure see Section 4.4). In a second step, estimated u_t and Σ_u can be used to identify ε_t by exploiting the below relationships following from 4.2 and 4.4:

$$\begin{aligned} y_t &= \sum_{i=0}^{\infty} \Phi_i u_{t-i} = \sum_{i=0}^{\infty} \Theta_i \varepsilon_{t-i} \\ \Phi_0 u_t &= \Theta_0 \varepsilon_t \\ u_t &= \Theta_0 \varepsilon_t \end{aligned}$$

where the last step follows from the fact that $\Phi_0 = I_k$. But then, it also follows that:

$$\begin{aligned} \Theta_0^{-1} u_t &= \varepsilon_t \\ \varepsilon_t &= A_0^* u_t \end{aligned}$$

as $\Theta_0^{-1} = A_0^*$ and $A_0^* = B \cdot A_0 = B$ since $A_0 = I_k$. Consequently, identification of ε_t only requires identification of B, which is nothing else than the coefficient matrix of the structural VAR(p) model. As such, it contains $(K \times K) = K^2$ elements. In order to identify K^2 elements, K^2 restrictions (i.e. equations) are required. By assuming $\Sigma_\varepsilon = I_K$, $(K^2 - K)/2$ restrictions have already been imposed. Thus, in order to have an identified system, one needs additional $(K^2 + K)/2$ restrictions.

In the main VARX(p, q) specification of this study, these additional restrictions are imposed by means of a recursive identification scheme achieved through a Choleski decomposition of the variance covariance matrix Σ_u (Kilian & Lütkepohl, 2017, pp.219-218). The next section will go into the details of such identification strategy.

4.2 Identification strategy

As a variance covariance matrix, Σ_u is real valued, symmetric, positive and definite. Therefore, it has the following Choleski decomposition:

$$\Sigma_u = PP'$$

where P is a lower triangular matrix with standard deviations of the error terms u_t as main diagonal elements (Lütkepohl, 2005, p.658). Then, the additional $(K^2 + K)/2$ necessary restrictions can be imposed by setting $\Sigma_u = PP'$ and $B^{-1} = P$. Consequently, exploiting the relationship $PP^{-1} = I_K$, it holds that:

$$\begin{aligned} y_t &= \Phi(L)PP^{-1}u_t \\ &= \Theta(L)\varepsilon_t \end{aligned}$$

where $\Theta(L) = \Phi(L) \cdot P = A(L)^{-1} \cdot P$ and $\varepsilon_t = P^{-1} \cdot u_t$. It can now be shown that, as assumed, $\Sigma_\varepsilon = I_K$, as:

$$\begin{aligned} \Sigma_\varepsilon &= E(\varepsilon_t \varepsilon_t') = E[(P^{-1}u_t)(P^{-1}u_t)'] = P^{-1}E(u_t u_t')P^{-1\prime} \\ &= P^{-1}\Sigma_u P^{-1\prime} = P^{-1}PP'P^{-1\prime} = I_K \end{aligned}$$

As P is lower triangular, so is B^{-1}. Therefore, setting the additional necessary restrictions by means of a Choleski decomposition implies imposing a certain ordering of the variables. In other words, as B is used to derive the structural representation of the model, y_{st} cannot have any instantaneous effect on y_{Kt} if $K < s$ (Lütkepohl, 2005, p.59). Economically, this means that some of the structural shocks in ε_t do not have a contemporaneous impact on some of the endogenous variables. Thus, a plausible and realistic ordering needs to be chosen based on economic interpretation and reasoning.

As the aim of the analysis is to estimate the effect caused by a shock in the exchange rate, the position of the exchange rate relative to the other variables is of most importance.

However, as mentioned by Hahn (2007) in a similar study, there are multiple plausible positions for the exchange rate. If one believes the exchange rate to quickly adjust to new information, such as changes in policy rates, a plausible choice would be to place it last. However, this would imply ex ante restricting the contemporaneous impact of exchange rate shocks to all other variables, including prices, to zero. In light of the transmission mechanism to be estimated here, this position is not applicable to this study. Alternatively, one could order it first. The advantage of doing so is that, ex ante, it does not require to assume zero contemporaneous impact on any of the variables. Economically, this might be justified by the idea that the exchange rate is influenced mainly by external developments and less by domestic variables, at least instantaneously, especially considering the substantial delay with which domestic economic data (such as GDP) is published (Hahn, 2007, p.12). On the other hand, as done by Peersman and Smets (2001), one could also argue that monetary shocks (i.e. shocks to the money supply and the exchange rate) affect the real economic activity only with a lag (p.9). Taking this as the most reasonable assumption, thereby also following Stulz (2007)'s ordering choice in a similar study on the Swiss economy, the exchange rate is positioned before import and consumer prices but after the output gap and the monetary policy variable. Formally:

$$y_t = (gdp_t, \Delta m_t, \Delta e_t, \Delta ipi_t, \Delta cpi_t)' \tag{4.5}$$

This ordering is in line with other relevant literature using VAR models to estimate ERPTs (see, for instance, Choudhri et al., 2005; Hahn, 2003; McCarthy, 2007) and is motivated by the following economic intuition: as already mentioned, it seems reasonable to assume that monetary shocks (i.e. shocks to the money supply and the exchange rate) affect the real economic activity only with a lag. Furthermore, the asset price nature of the exchange rate makes it react immediately to shocks to real and monetary variables. With regards to prices, positioning the exchange rate before the import and consumer price indexes reflects the idea that prices are set along the distribution chain and that the pass-through to consumer prices involves two stages: the first on the import prices and the second on the consumer prices.

4.3 Impulse response functions derivation

Once the economically meaningful innovations ε_t have been identified, IRF can be derived. As $\Sigma_\varepsilon = I_K$, the elements of ε_t are uncorrelated. This means that it is possible to analyse the effect of a shock in one variable at a time while keeping the others constant (Lütkepohl, 2005, p.58). Based on Lütkepohl (2005) and the theoretical considerations made above, the orthogonalised impulse responses are contained in the MA coefficient matrices Θ_i, where $\Theta_i = \Phi_i \cdot P$. Furthermore, it is known that $\Phi_i = J \cdot A^i \cdot J'$, where J is a $(K \times Kp)$ matrix of the form $[I_K : 0 : ... : 0]$[3] (pp.18, 57-62). Therefore, orthogonalised IRF for i periods after the shock can be computed recursively in the following way:

$$\begin{aligned}
\Theta_0 &= \Phi_0 P = I_k P \\
\Theta_1 &= \Phi_1 P = JAJ'P \\
\Theta_2 &= \Phi_2 P = JA^2J'P \\
&\vdots \\
\Theta_i &= \Phi_i P = JA^iJ'P
\end{aligned}$$

In order to compute IRF as pass-through elasticities, i.e. as percentage changes in prices after a one percentage point increase in the exchange rate or in import prices, the shocks need to be normalised to one-unit shocks, as variables are included in first difference of their natural logarithm. To do this, the Cholesky lower matrix P is scaled with the inverse of its main diagonal, D^{-1}, so that $\tilde{P} = P \cdot D^{-1}$ is used instead of P. Then, the elements of Θ_i represent responses of the system to such one-unit innovations.

As commonly done in the literature (see, for instance, Stulz, 2007), when estimating pass-through elasticities, accumulated IRF are the most informative result, as one is interested in the cumulative percentage change in prices over time after a shock. These are defined as:

$$\Psi_i = \sum_{i=0}^{n} \Theta_i$$

where $n = 1, ..., 24$, i.e. this study looks at the effects of a shock over 24 months. As the most informative, graphically, results are shown in terms of accumulated IRF only.

4.4 Estimation method

The baseline VARX(p, q) model is estimated by means of multivariate least square (MLS) estimation. Again, to simplify notation, the exogenous variable x_t will be ignored in the derivations to follow, as it can easily be added in each step. In a low dimensional setting, i.e. where the number of parameters to be estimated is substantially lower than the length of the time series, the model can be estimated by MLS even in the presence of exogenous variables (Nicholson et al., 2017, p.629)[4]. Following (Lütkepohl, 2005, pp. 69-77), for the purpose of estimation, it is useful to rewrite the model in the following

[3] As explained by Lütkepohl (2005), every VAR(p) process can be written in a VAR(1) form. Matrix J is introduced in the context of switching from the VAR(1) form of a VAR(p) process to its traditional representation, such as 4.1.

[4] See also the work of Ocampo and Rodriguez (2012) on this.

notation:

$$\begin{aligned}
Y &= [y_1, ..., y_T] && (K \times T)\\
B &= [\nu, A_1, ..., A_p] && (K \times (Kp+1))\\
Z_t &= [1, y_{t-1}, ..., y_{t-p}]' && ((Kp+1) \times 1)\\
Z &= [Z_0, Z_1, ..., Z_T] && ((Kp+1) \times T)\\
U &= [u_1, ..., u_T] && (K \times T)
\end{aligned}$$

Then, the model can be written as:

$$Y = BZ + U$$

and the MLS estimator can be derived as the conventional OLS estimator for B:

$$\hat{B} = YZ'(ZZ')^{-1}$$

Residuals and their variance covariance matrix can be estimated from $\hat{B}$:

$$\begin{aligned}
\hat{U} &= Y - \hat{B}Z\\
\hat{\Sigma}_u &= \frac{1}{T}(Y - \hat{B}Z)(Y - \hat{B}Z)'
\end{aligned}$$

IRF and respective confidence intervals can then be estimated using $\hat{\Sigma}_u$ and $\hat{B}$. Details on the computation of the confidence intervals are included in Appendix B.

4.5 Model specification

In the previous sections, lag orders p and q were assumed to be known. However, they both need to be chosen on the basis of the available sample data. The selection of p and q is important as choosing them unnecessarily large will reduce the precision of the model and of the impulse reponses (Lütkepohl, 2005, p.135). Therefore, it is of relevance to have procedures and criteria to make the appropriate choice.

As highlighted by Lütkepohl (2005) "because different criteria emphasize different aspects of the data generation process and may therefore all provide useful information for the analyst, it is common not to rely on just one procedure or criterion for model choice but use a number of different statistical tools." (p.157). Following this practice, the lag order is selected based on multiple criteria, namely a top-down sequential Likelihood Ratio (LR) testing procedure, Akaike's Final Prediction Error (FPE), AIC, Hannah-Quinn criterion (HQ) and Schwarz Criterion (SC). As white noise residuals are always a good indication for the fitted model well describing the data generation process (DGP), to complement the lag order selection procedure, a final Portmanteau test for residual autocorrelation is implemented on residuals of some model candidates.

Test results as well as computational details on the testing procedures are presented in Appendix C. Note that tests have been implemented by setting $p = q$. As shown by Table C.1 in Appendix C , when the maximal lag is set to 4, despite the sequential LR test not being able to select an order, FPE and AIC select 4 lags while HQ and SC select 1. When the maximal lag is increased to 8 or 12, the FPE and AIC select 6 lags and HQ and SC again 1.

Given the use of monthly data, the inclusion of just one lag seems too parsimonious. Three lags would represent an order in between those selected by the different criteria. Additionally, from an economic perspective, three lags would represent a quarter. Therefore, Portmanteau tests have been carried out on model specifications with 3, 4 and 6 lags. For all three, the null hypothesis of white noise residuals could not be rejected at the 1% significance level. In light of these results, the most parsimonious model amongst those with more than one lag has been chosen, namely with 3 lags for both endogenous and exogenous variables[5]. Furthermore, a verification of the stability condition[6] shows that the chosen model is stable.

[5] Estimations with 4 or 6 lags show that results change little compared to the chosen order of 3 lags (see Figure C.1 in Appendix C).

[6] A VAR(p) process is stable if its reverse characteristic polynomial has no roots in and on the complex unit circle, i.e. if $\det(I_K - A_1 z - ... - A_p z^p) \neq 0$ for $|z| \leq 1$ (Lütkepohl, 2005, p.16).

5 Results

The baseline specification treats the entire sample period as a single economic regime characterised by the same VARX(p, q) process. It therefore assumes that the transmission mechanism of exchange rate and import price changes to import and consumer prices remains unchanged over the considered time period 2000:01 to 2019:09. The extent to which this assumption is likely to hold will be investigated in a second step of the analysis through a series of structural break Chow tests and the estimation of an intervention model (see Section 7).

This section presents pass-through elasticities measured through the computation of IRF based on the VARX(3,3) model specified in the previous section by treating the entire sample period as a single economic regime. More precisely, the system is shocked by a structural innovation in the equation describing the exchange rate or import prices. Responses to these shocks are then depicted for import prices and/or consumer prices. Impulse responses are computed for 24 months after the shock is sent into the system, so that pass-through patterns are observed over two years after a shock. As shocks have been normalised to one-unit, impulse responses represent the percentage change in import or consumer prices after a one percentage point increase in the exchange rate or import prices[1]. Accumulated responses show the cumulative percentage change over all previous and current periods.

5.1 Exchange rate pass-through to import prices

Panel (a) of Figure 5.1 depicts accumulated impulse responses of import prices to a one-unit shock in the exchange rate. First of all, as expected, import prices increase following a depreciation of the Swiss Franc against the Euro. Secondly, their reaction seems to be quite fast: already after three months, the pass-through amounts to 0.22 and its peak of 0.50 is reached after nine months (see second column of Table 5.1). In the long-run, i.e. after two years, the pass-through amounts to 0.42. Overall, results suggest an incomplete but significantly different from zero ERPT to import prices. Confidence bounds always different from one and zero confirm this.

Results are in line with related literature on Switzerland. Stulz (2007), who also implements a VAR analysis and computes pass-through elasticities, finds an incomplete but significant pass-through of 0.37 after two years. The difference could be mainly due

[1] As the nominal exchange rate is inserted as Swiss Francs over Euros, an increase implies a depreciation of the Swiss Franc.

to two facts. First, Stulz (2007) looks at a different sample period, namely from 1976 to 2004, when no drastic exchange rate appreciation or monetary policy intervention had happened yet. Second, instead of focusing only on the *EURCHF* nominal exchange rate, his analysis focuses on the nominal effective exchange rate of the Swiss Franc versus 24 trading partners. Žídek and Šuterová (2017), who also look at a general effective exchange rate but at a similar period to this study, namely 2000 to 2016, find a higher pass-through of 0.67 percent two years after a one percent increase in the exchange rate. Similarly, Choudhri and Hakura (2015)'s VAR estimation of the Swiss ERPT elasticity on import prices amounts to 0.52. On the other hand, for Switzerland, Campa and Goldberg (2005) find an almost complete long-run pass-through elasticity on import prices of 0.93. However, other than the difference in the exchange rate considered (since, similarly as in the other study, authors look at the effective nominal exchang rate) the discrepancy with the results of this study could be due to differences in methods applied: Campa and Goldberg (2005) compute the pass-through elasticity as the coefficient of a linear regression using quarterly data.

5.2 Pass-through of import prices to consumer prices

Panel (b) of Figure 5.1 depicts accumulated impulse responses of consumer prices to a one-unit shock in import prices. As for the ERPT, the import price pass-through to consumer prices is incomplete but significantly different from zero, as confidence bounds confirm. Also similarly to the ERPT to import prices, the highest pass-through is reached after nine months, when it amounts to approximately 0.27 (see the fourth column of Table 5.1). On the other hand, import price pass-through to consumer prices remains much more constant over the two years after the shock than the ERPT to import prices. Indeed, in the long-run, it slightly decreases to 0.25. As highlighted in theoretical considerations, the pass-through to consumer prices, i.e the second channel through which exchange rate changes affect consumer prices[2], strongly depends on the import penetration of the economy. According to the SFSO, of the goods included in the Swiss CPI in 2019, only 25.4 percent were imported goods. Moreover, in the last years, this share has remained quite constant[3]. The share of imported goods within the basket of consumption goods is extremely important when interpreting the transmission mechanism from import prices to consumer prices. As well explained by Stulz (2007), ignoring potential direct effects of foreign shocks on prices of domestic goods competing against imported ones[4], if prices

[2] As explained in Section 4, the first channel is represented by changes in imported goods, which are part of the basket of consumer goods.

[3] Source: SFSO's weighting of the CPI (Landesindex der Konsumentenpreise - Gewichtung) for the years from 2011 to 2019.

[4] For instance, domestic producers might adjust their prices in reaction to changes in prices of competing imported goods, especially if these are considered strong substitutes.

were fully passed through along the distribution chain, the pass-through from import prices to consumer prices should equal the share of imported goods (p.13). In other words, if the import prices pass-through to consumer prices equals this share, it might be considered complete. In light of this, the import prices pass-through estimated here is surprisingly strong. As can be seen in the fourth column of Table 5.1, the accumulated impulse responses are always very close to the mentioned share of imported goods. After three months, they even surpass it. This, however, is not completely unlikely if direct effects on domestic producers exist. For instance, domestic producers might decide to increase prices in reaction to an increase in prices of imported goods if these represent good substitutes.

5.3 Exchange rate pass-through to consumer prices

Panel (c) of Figure 5.1 depicts accumulated impulse responses of consumer prices to a one-unit shock in the exchange rate. In other words, this could be interpreted as a reduced form estimation of the ERPT. A depreciation of the Swiss Franc caused by a one percentage point increase in the *EURCHF* exchange rate increases consumer prices by about 0.04 percent after 3 months, 0.07 after nine and, in the long run, by only 0.06 (see the last column of Table 5.1). Therefore, results and confidence bounds indicate that the direct pass-through to consumer prices is significantly different from zero, at least after three months, but weak, especially if compared with the pass-through to import prices. This, however, is quite in line with existing empirical results for Switzerland[5]. More generally, many studies have provided empirical evidence for a decreasing ERPT along the distribution chain. To mention a few, Chung et al. (2011) reach the same conclusion when looking at Australia, Cheikh and Louhichi (2015) and Hahn (2003) when looking at countries in the Euro Area.

As only a share of the basket of consumer goods is supposed to be directly affected by exchange rate changes, namely the share of imported goods, these results are not surprising. In fact, to get a sense of the expected magnitude of these pass-through elasticities, one should compare them with the product of the ERPT on import prices with the pass-through of import prices to consumer prices. Doing so reveals that the ERPT to consumer prices is still lower than this product, but not too far from it.

Overall, estimations over the entire sample period suggest an incomplete but substantial ERPT to import prices and a very strong transmission mechanism from import prices to consumer prices. Taken together, this seems to indicate that the low ERPT to consumer prices is likely to be primarily blocked by sticky import prices (i.e. by the incomplete

[5] Stulz (2007) found an ERPT of 0.09 after three months and of 0.18 after two years. Žídek and Šuterová (2017) found a pass-through of 0.12 percent after two years.

ERPT to import prices) and a low share of imported goods among the consumer goods. Reasons behind import price stickiness could be exporting firms wanting to keep their competitiveness, and, as mentioned earlier, the fact that importers often add factors of production such as transport and labor so that final import prices are not necessarily fully affected by exchange rate changes.

With that said, in light of the substantial appreciation experienced by the Swiss Franc against the Euro during the period considered, one might still expect a stronger direct effect of exchange rate changes on consumer prices. Given the substantial pass-through to import prices, such an appreciation results in relatively cheaper imports to Switzerland from the Eurozone. Consequently, in theory, one would expect the share of imported goods in the basket of consumer goods to increase, thereby increasing also the measured pass-through to consumer prices. However, as reported by the SFSO and as already mentioned above, the share of imported goods amounts to only 25.4 percent in 2019. 74.6 percent of the composition of the CPI is made out of goods priced in the domestic currency and thus not directly affected by exchange rate changes. Furthermore, even assuming strong direct effects on domestic goods due to high substitutability with imported ones, 59.6 percent of the basket of the CPI would still represent non-tradable services and so remain rather unaffected to exchange rate changes. Moreover, these numbers have changed little during the sample period considered. Even at times when the appreciation reached extremely high levels, such as in 2010, the share of imported goods amounted to only 27.0 percent[6]. In light of this, low estimates of ERPT to consumer prices seem reasonable.

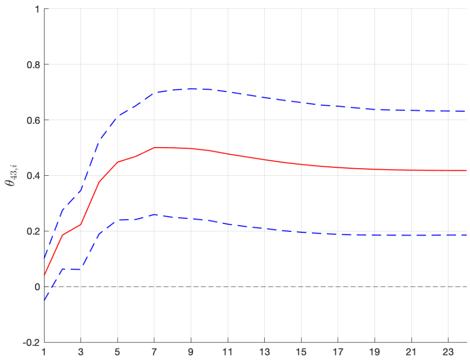
Table 5.1: Impulse Responses and Accumulated Impulse Responses

	Δe_t to Δipi_t		Δipi_t to Δcpi_t		Δe_t to Δcpi_t	
Period i	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**
1	0.039279	0.039279	0.180346	0.180346	0.006852	0.006852
3	0.038015	0.223463	0.026027	0.247098	0.012311	0.041553
6	0.020968	0.468933	0.004262	0.271149	0.003040	0.067291
9	-0.002651	0.497144	-0.002379	0.265205	-0.000312	0.073161
12	-0.010349	0.467162	-0.002860	0.255442	-0.001496	0.069108
15	-0.007554	0.439763	-0.001596	0.249313	-0.001069	0.065310
18	-0.003647	0.424909	-0.000626	0.246595	-0.000529	0.063173
21	-0.001211	0.419203	-0.000147	0.245777	-0.000180	0.062332
24	-0.000168	0.417899	0.000017	0.245712	-0.000028	0.062129

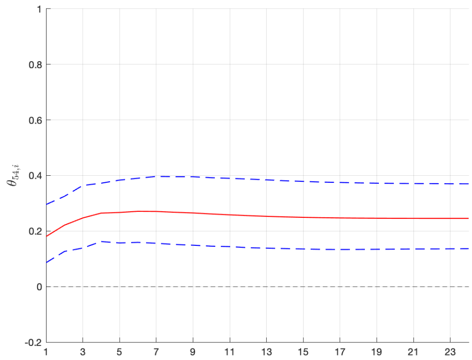
Notes: The table depicts orthogonalised impulse responses (IRF) and accumulated orthogonalised impulse responses (Acc. IRF) for a one-unit shock in the impulse variable (Δe_t or Δipi_t) based on the specified VARX(3,3) model. Period i refers to the number of months after the realisation of the shock.

[6] Source: SFSO's weighting of the CPI for 2010 (Landesindex der Konsumentenpreise - Gewichtung 2010).

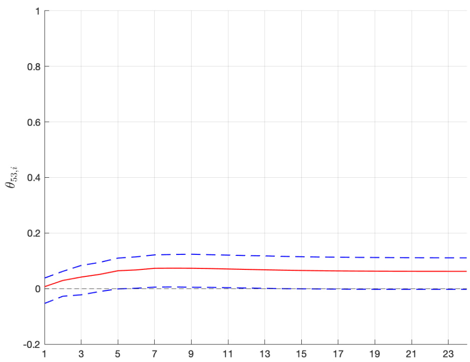
Figure 5.1: Accumulated Impulse Response Functions



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices



(c) ERPT to Consumer Prices

Notes: The figure shows accumulated orthogonalised impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a one-unit shock in the impulse variable based on the specified VARX(3,3) model. Blue dashed lines represent 95% standard residual-based recursive design bootstrap confidence bounds.

6 Sensitivity analysis

Estimated pass-through elasticities have been derived through a VARX(3,3) model by making a number of choices and assumptions. This section investigates the robustness of results with respect to such choices and assumptions. More precisely, it aims to shed light on the sensitivity of the estimates to different Choleski orderings, to changes in the monetary policy variable and, finally, to the choice of addressing nonstationarity issues with first differencing instead of a VECM.

6.1 Alternative Choleski orderings and generalized impulse response functions

Changing the ordering of the variables gives insights on the robustness of results as impulse responses may be highly sensitive to and determined by it (Lütkepohl, 2005, p.61). The importance of the ordering depends on the magnitude of the correlation coefficients between the elements of the residuals of the reduced form, u_t (Enders, 2010, p.298). The correlation coefficients of the residuals are given by:

$$\rho_u = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ -0.614 & 1 & 0 & 0 & 0 \\ 0.253 & -0.426 & 1 & 0 & 0 \\ 0.093 & -0.559 & 0.048 & 1 & 0 \\ -0.076 & -0.634 & -0.047 & 0.875 & 1 \end{pmatrix}$$

where the zero entries indicate that the contemporaneous correlations for those directions were restricted to zero when imposing the baseline ordering. Surprisingly, the correlations between u_{e_t} and u_{ipi_t} and u_{e_t} and u_{ipi_t} are rather low. However, others are quite high, such as between u_{gdp_t} and u_{m_t}, u_{gdp_t} and u_{e_t} or u_{ipi_t} and u_{m_t}. This highlights the importance of computing estimates also with different orderings.

As there are five endogenous variables, there are $5! = 120$ possible orderings. Nevertheless, some assumptions can be made in order to limit the set of plausible orderings. In the baseline specification, three assumptions were made with regards to this. First, it has been assumed that prices are set along the distribution chain, such that import prices are ordered before consumer prices. Second, as the effects of exchange rate changes on prices are of main interest, the former have been ordered before the latter, assuming that prices affect the exchange rate only with a lag. Lastly, real economic activity (i.e.

gdp_t) has been assumed to react only with a lag to monetary policy, so that the output gap was ordered before the monetary policy variable. Even while maintaining all three assumptions, other four potential orderings can be imposed. These are represented by 6.1, 6.2, 6.3 and 6.4 below. What changes in these orderings compared to the baseline specification is the position of Δe_t relative to Δm_t and gdp_t. It is no longer assumed that it immediately reacts to both real economy and monetary shocks. On the contrary, it is either ordered first to reflect the idea highlighted by Hahn (2007) that it might be mainly influenced by external developments rather than domestic variables (see 6.3 and 6.4), or it is ordered before the monetary policy variable, so to allow monetary policy to contemporaneously react to shocks in the exchange rate (as done by McCarthy (2007)) (see 6.1 and 6.2).

$$y_t = (gdp_t, \Delta e_t, \Delta m_t, \Delta ipi_t, \Delta cpi_t)' \tag{6.1}$$

$$y_t = (gdp_t, \Delta e_t, \Delta ipi_t, \Delta m_t, \Delta cpi_t)' \tag{6.2}$$

$$y_t = (\Delta e_t, gdp_t, \Delta m_t, \Delta ipi_t, \Delta cpi_t)' \tag{6.3}$$

$$y_t = (\Delta e_t, gdp_t, \Delta ipi_t, \Delta m_t, \Delta cpi_t)' \tag{6.4}$$

$$y_t = (\Delta m_t, \Delta e_t, gdp_t, \Delta ipi_t, \Delta cpi_t)' \tag{6.5}$$

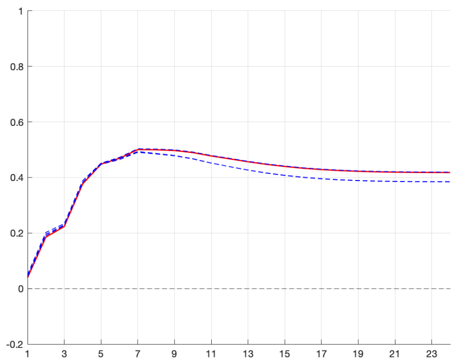
$$y_t = (\Delta m_t, \Delta e_t, \Delta ipi_t, gdp_t, \Delta cpi_t)' \tag{6.6}$$

Relaxing the third assumption of the baseline specification, one can test other two potential orderings, represented by 6.5 and 6.6. Here, real economic activity is allowed to contemporaneously react to monetary policy and exchange rate shocks.

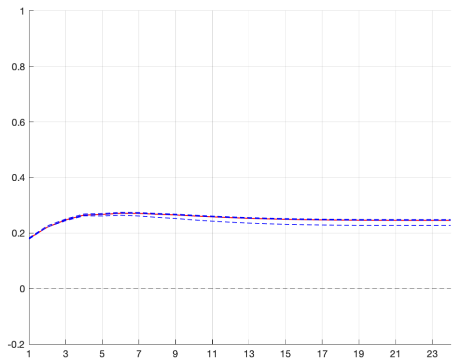
Figure 6.1 compares the IRF of the baseline specification with the six alternative orderings described above. The plots show that results change little when using one of the alternative orderings. The ERPT to import prices (panel a) seems to decline slightly more and faster after the seventh period for some alternative orderings, but the shape and the magnitude of the IRF remain almost unchanged. Even more robust seem to be the estimations of the pass-through of import prices to consumer prices and the ERPT to consumer prices (panels b and c), as the accumulated impulse responses are even closer to each other.

To sum up, the shape and the magnitude of the IRF change very little when using different Choleski orderings. This suggests that, despite the components of the error terms being highly correlated, results are quite robust to different orderings.

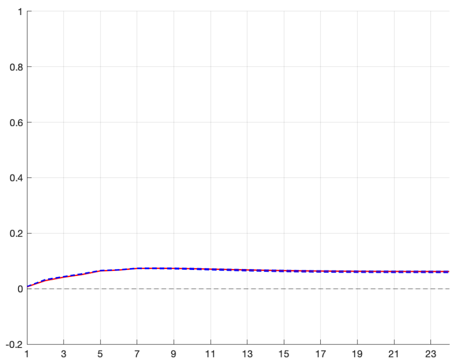
Figure 6.1: Accumulated Impulse Response Functions under Different Orderings



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices



(c) ERPT to Consumer Prices

Notes: The figure shows accumulated orthogonalised impulse response functions under different orderings. Red solid lines represent orthogonalised accumulated impulse responses to a one-unit shock in the impulse variable computed with the baseline specification ordering, based on the specified VARX(3,3) model. The blue dashed lines represent the same accumulated orthogonalised impulse responses, but based on six different alternative orderings.

Another common way of checking the robustness of results to an imposed ordering is to estimated generalized impulse response functions (GIRF). Pesaran and Shin (1998) proposed this approach to compute unique impulse responses invariant to the ordering of the variables in the system. As they explain, GIRF measure the effect of one standard error shock to the j^{th} equation at time t on expected values of y at time $t+i$, where i represents the number of periods after the shock has been sent. They are computed as:

$$\psi_j(i) = \sigma_{jj}^{-\frac{1}{2}} A_i \Sigma_u e_j$$

where e_j is a $(K \times 1)$ vector with unity as its j^{th} element and zeros elsewhere, A_i corresponds to $J \cdot A^i \cdot J'$ (as computed in Section 4.3 using coefficient matrices of the reduced form) and Σ_u is the variance covariance matrix of error terms u_t (p.19)[1]. Estimated ac-

[1] For further conceptual and computational details on GIRF, see Pesaran and Shin (1998).

cumulated GIRF are depicted in Figure D.1 in Appendix D, where they are compared to accumulated IRF based on the baseline ordering. Pass-through estimates are broadly robust, especially those of ERPT to import prices and ERPT to consumer prices. Using GIRF instead of orthogonalized IRF does not change the broad pattern and magnitude of the transmission of exchange rate shocks to import and consumer prices. This is taken as a further indication for the baseline specification being robust to different orderings. Most importantly, they do not seem to affect impulse responses to an important extent.

6.2 Changing the monetary policy variable

As mentioned in Section 3.1, in order to control for monetary policy actions when estimating ERPTs via VAR models, the most common variable used is a short-term interest rate (see, for instance, An & Wang, 2012; Hahn, 2003; McCarthy, 2007). Due to the Swiss monetary policy context of the period considered here (see Section 3.1 for a more detailed explanation), in its baseline specification, this study deviates from standard literature by including the monetary aggregate M3 instead. Nevertheless, in light of the relevant role of short-term interest rates for monetary policy, it is of interest to check the sensitivity of results to this choice by reestimating them with some short-term interest rate as monetary policy variable.

Fink et al. (2020) and Ranaldo and Rossi (2010) all find that the SNB policy rate[2] significantly affects the Swiss Franc in that unexpected policy rate hikes lead to its appreciation. Furthermore, they find that unexpected policy rate changes affect the yield curve of Swiss government bonds, especially that of bonds with 2-years maturity. Based on these findings, to check the sensitivity of results to the choice of the monetary policy variable, IRF have been recomputed using the SNB policy rate, the yield on the 1-year maturity Swiss government bond and the yield on the 2-years maturity Swiss government bond[3]. Interest rate measures have not been transformed in their logarithmic form. However, as for the monetary aggregate M3, they have been included in first differences.

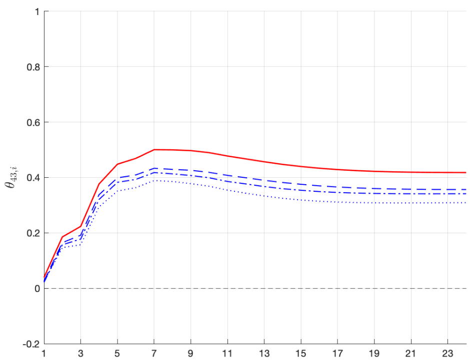
Results are depicted in Figure 6.2. As can be seen by the lines being extremely close to each other, with regards to the pass-through from import to consumer prices (panel b) and ERPT to consumer prices (panel c), results are almost identical to the baseline specification. Slightly larger differences can be found when looking at the ERPT to import prices (panel a): for all interest rate measures, the pass-through seems to be slightly smaller than that computed with the monetary aggregate M3. However, differences are

[2] Until June 2019, the SNB policy rate has been the 3-month Libor, i.e. the Libor for 3-month interbank loans in Swiss Francs. The SNB used to set a range and, as a rule, aimed to keep the policy rate in the middle of that range. In June 2019, the SNB decided to change the policy rate to the SARON (Swiss Average Rate Overnight).

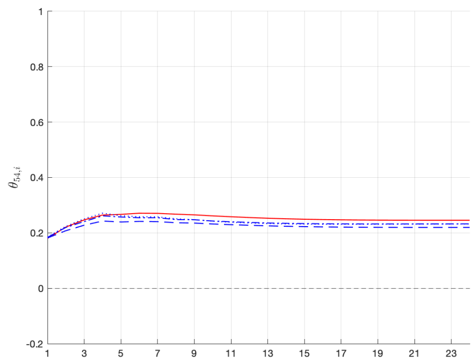
[3] Data on these rates has been retrieved from the SNB database. For the SNB policy rate, the middle of the range has been used.

small and the shape of the responses remains unchanged. This suggests that results are quite robust to the choice of the monetary policy variable.

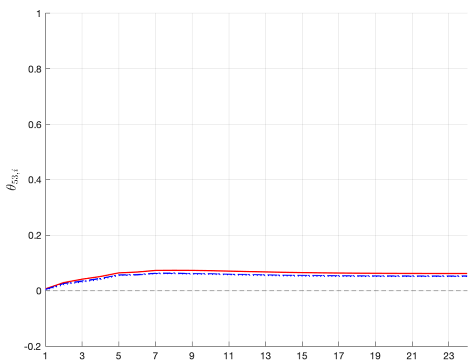
Figure 6.2: Accumulated Impulse Response Functions under Different Monetary Policy Variables



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices



(c) ERPT to Consumer Prices

Notes: The figure shows accumulated orthogonalised impulse response functions under different monetary policy variables. Red solid lines represent baseline specification accumulated orthogonalised impulse responses based on the monetary policy variable M3. Blue long dashed lines, blue dash-dotted lines and blue dotted lines represent accumulated orthogonalised impulse responses based on the SNB policy rate, the 1-year yield and 2-year yield of Swiss government bonds as monetary policy variable respectively.

6.3 A vector error correction model

Often economic variables exhibit upward or downward movements, a feature which can be generated by stochastic trends in integrated variables. If the same stochastic trend is driving a set of integrated variables jointly, they are said to be cointegrated. Then, stationary linear combinations of cointegrated variables are called cointegrating relationships and can be seen as long-term equilibrium relationships between the cointegrated economic variables (Kilian & Lütkepohl, 2017, p.75). Cointegrating relationships can be imposed by reparameterizing the VAR model in levels as a VECM, which offers the great feature of analyzing the original variables (or their logarithms) rather than the rates of change while still accomodating data nonstationarity (Lütkepohl, 2005, p.237).

As explained by Kilian and Lütkepohl (2017) and briefly mentioned in Section 4.1, in the presence of unit roots, it is rarely clear when to use a VECM as opposed to a VAR model in levels or first differences. Nevertheless, if the integrated variables are not cointegrated, it is still preferable to estimate a VAR in first differences (Enders, 2010, p.384). Given the strong evidence suggesting the presence of unit roots in the series used, it seems reasonable to to estimate a VECM as an alternative specification.
Consider again a K-dimensional VAR(p) model in (log-)levels and reduced form:

$$y_t = \nu + A_1 y_{t-1} + ... + A_p y_{t-p} + u_t$$

where u_t is a standard white noise process. It is assumed that first differences $\Delta y_t = y_t - y_{t-1}$ are stationary and also that $det(I_K - A_1 z - ... - A_p z^p)$ has all its roots outside of the complex unit circle except for possibly some of them being equal to 1. In other words, the system is allowed to be nonstationary. Then, the $(K \times K)$ matrix

$$\Pi = -(I_k - A_1 - ... - A_p)$$

is singular with rank $r \leq K$ and can be expressed as the product of a $(K \times r)$ matrix α and a $(r \times K)$ matrix β, both with rank r. That is $\Pi = \alpha\beta$. β is the so called cointegrating matrix such that βy_t is stationary and represents the long-term equilibrium relationship between the economic variables. α is referred to as the loading matrix and indicates the speed of adjustment towards the long-term equilibrium. Rank r of matrix Π is of utmost relevance: if $r = 0$, Δy_t has a stable VAR($p-1$) representation, if $r = K$, $|I_K - A_1 - ... - A_p| = |-\Pi| \neq 0$ and, hence, the VAR operator has no unit roots so that y_t is a stable VAR(p) process. If, however, $0 < r < K$, then elements of y_t are cointegrated and have r cointegrating relationships (Lütkepohl, 2005, pp.245-249). In this case, one would like to reparametize the above VAR(p) model so that it contains the long-term relationship $\Pi = \alpha\beta$. This can be done by rewriting it as a VECM($p-1$) process of the following form:

$$\Delta y_t = \Pi y_{t-1} + \Gamma_1 \Delta y_{t-1} + ... + \Gamma_{p-1} \Delta y_{t-p+1} + u_t \tag{6.7}$$

where $\Gamma_i = -(A_{i+1} + ... + A_p)$ with $i = 1, ..., p-1$. Depending on whether variables exhibit a trend or not, both an intercept term and linear time trend term can be added either by leaving them unrestricted or by making sure they are absorbed in the cointegrating relationship, in which case they are said to be restricted (Kilian & Lütkepohl, 2017, p.81).

In what follows, in order to investigate on the existence of cointegrating relationships between the variables used and on the specification of a potential VECM, the Johansen methodology is applied by first estimating the rank r of matrix Π and then, secondly, a

VECM in accordance (see Johansen (1995) and Lütkepohl (2005)).

6.3.1 Testing for the rank of cointegration

Following Johansen (1995) and Lütkepohl (2005), in order to test whether there exist cointegrating relationships between the variables used, a VAR(p) model is estimated in log-levels. In a second step, once matrix Π is identified, Johansen tests are used to assess whether or not cointegration exists in the system of variables.

Before proceeding with estimations, it is worth mentioning a few words about the inclusion of the HCPI (in logs) ($hcpi_t$), which was modelled as exogenous in the baseline VAR specification in first differences. When cointegrating relationships are introduced, the question about how to split the contributions of this regressor between the cointegrating relationship and the short-term dynamics part of the VECM system arises. In light of the fact that the HCPI is likely to be correlated with the other variables (reason for which it was considered in the first place), and, therefore, might be cointegrated with them, in the context of a VECM, it seams reasonable to include it as an endogenous variable. This also allows it to potentially be in a long-term relationship with the other variables.

Lag order selection for the VAR(p) model to be estimated in log-levels has been based on the same criteria used for the baseline specification (see Table C.2 in Appendix C) in combination with well behaved residuals. As suggested by the sequential LR test, lag length has been set to two. Despite all the other criterias suggesting one lag instead, residual analysis suggests to include two. Furthermore, it seems reasonable and desirable in light of the analysis to include short-term dynamics terms in the VECM.

Once matrix Π has been identified by estimating the VAR(2) model in log-levels, one can proceed with testing its rank. To this end, following Kilian and Lütkepohl (2017), the sequence of hypotheses below may be considered:

$$H_0(r_0)\text{: rank }(\Pi) = r_0 \text{ versus } H_1(r_0)\text{: rank }(\Pi) > r_0$$

Then, under suitable regularity conditions, the corresponding LR test statistic is the so called trace statistic:

$$\lambda^{trace}(r_0) = -2(\log\, l(r_0) - \log\, l(K)) = -T \sum_{i=r_0+1}^{K} \log(1-\lambda_i)$$

where $l(r)$ denotes the maximum of the likelihood function for the VECM($p-1$) model given cointegration rank r and λ_i the i^{th} eigenvalue of the $\hat{\Pi}$ matrix. Then, the rank is chosen to be the one of the first null hypothesis that cannot be rejected. The trace test is often performed in combination with the so called maximum eigenvalue test, which tests following alternative sequence of hypotheses:

$$H_0(r_0)\text{: rank }(\Pi) = r_0 \text{ versus } H_1(r_0+1)\text{: rank }(\Pi) = r_0 + 1$$

for $r_0 = 0, ..., K-1$. The corresponding LR test statistic is the following:

$$\lambda^{max}(r_0) = -T \log(1 - \lambda_{r_0+1})$$

As for the trace test, the testing procedure ends when the null hypothesis cannot be rejected for the first time (pp.100-101).

Table 6.1 shows results of trace and maximum eigenvalue tests based on maximum likelihood (ML) estimation of a VECM(1) model (in accordance with two lags selected for the VAR process). Since some of the variables included in the system exhibit a linear time trend, the model has been estimated with an unrestricted intercept.

As can be seen by comparing critical values with the test values, all tests reject rank

Table 6.1: Trace and Maximum Eigenvalue Tests for Cointegration Rank

			Critical values				**Critical values**		
λ_i	H_0	λ^{trace}	**10%**	**5%**	**1%**	λ^{max}	**10%**	**5%**	**1%**
0.2086	$r = 0$	134.9479	90.39	85.18	104.20	54.9815	36.35	39.43	51.30
0.1511	$r = 1$	79.9664	66.49	70.60	78.87	38.4853	30.84	33.32	35.80
0.0915	$r = 2$	41.4811	45.23	48.28	55.43	22.5556	24.78	27.14	29.16
0.0397	$r = 3$	18.9255	28.71	31.52	37.22	9.5096	18.90	21.07	22.89
0.0261	$r = 4$	9.4159	15.66	17.95	23.52	6.2225	12.91	14.90	17.07
0.0135	$r = 5$	3.1933	6.50	8.18	11.65	3.1933	6.50	8.18	11.65

Notes: The table shows test statistic and critical values of LR trace (λ^{trace})and maximum eigenvalue (λ^{max}) tests for cointegration rank. The tests are based on a VEC(1) model with unrestricted intercept. Critical values are shown for significance levels 10, 5 and 1% and are taken from Table 1.1 of Osterwald-Lenum (1992).

$r = 0$. Rank $r = 1$ is also rejected by both tests at all significance levels. The first null hypothesis that cannot be rejected is the same for trace and maximum eigenvalue tests and corresponds to rank $r = 2$. Therefore, the preferred cointegrating rank is $r = 2$ and, as alternative specification, a VECM(1) with $r = 2$ is estimated.

6.3.2 Results of the model

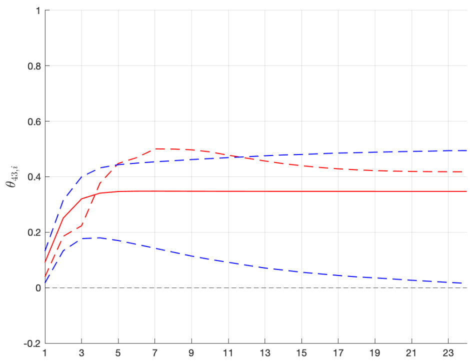
As alternative specification, based on test results derived in the previous section, a VECM(1) with rank $r = 2$ has been estimated by ML.

In order to compare it with the baseline specification, IRF need to be computed. As explained by Lütkepohl (2018), despite an unstable, integrated or cointegrated VAR(p) process not possessing a valid MA representation, impulse response analysis can be performed as for stationary processes (see Section 4.3) by retrieving the coefficient matrices of the VAR(p) process from the corresponding estimated VECM($p-1$). The difference, however, lies in the fact that, for stable processes, the responses ϕ_{ik} or θ_{ik} go to zero as $i \to \infty$, i.e the marginal response to an impulse in a stationary process is transitory. In contrast, in cointegrated systems, impulses can have permanent effects. That is, in a

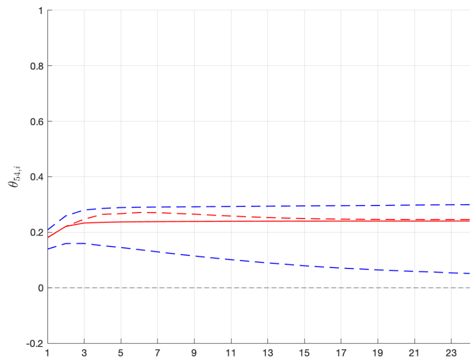
K-dimensional system with $r < K$ cointegrating relationships, at least $K - r$ of the K possible shocks (i.e. structural innovations) have permanent effects and at most r shocks have transitory effects (p.3). For the case considered here, this implies that at least 3 structural innovations have permanent effects on the system. Indeed, a remarkable feature often found in VECM impulse responses is that they do not die out to zero but rahter approach some nonzero value.

To identify the shocks and compute IRF, as in the baseline specification, a Choleski

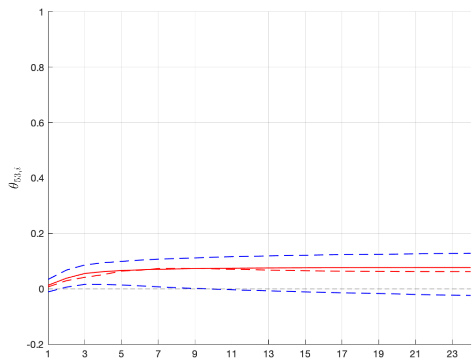
Figure 6.3: Impulse Responses of the Vector Error Correction Model



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices



(c) ERPT to Consumer Prices

Notes: The figure shows impulse response function of the VARX and the VEC models. Red solid lines represent orthogonalised impulse responses based on the specified VECM(1) model with rank 2. Red dashed lines represent the accumulated orthogonalised impulse response functions based on the baseline specification model VARX(3,3). Blue dashed lines represent 95% standard residual-based recursive design bootstrap confidence bounds for the impulse responses of the VECM(1).

decomposition is applied to the variance covariance matrix of the VECM(1) error terms u_t[4]. As expected, estimated orthogonalized impulse responses do not decay to zero but approach a positive constant approximately 3 months after the impulse is sent into the system, suggesting that shocks in both the exchange rate and import prices are permanent

[4] The same variable ordering as in the baseline specification has been imposed.

ones. This reflects the nonstationarity of the VAR(p) system. Consequently, in contrast to the baseline estimation, accumulated responses do not show any stabilization towards a certain constant. In light of this, it seems reasonable to compare impulse responses of the alternative specification with the accumulated ones of the baseline model. Figure 6.3 shows this comparison. The estimated reaction of consumer prices to a one unit shock in the exchange rate (panel b) or in import prices (panel c) seems to be extremely similar between the two models. Despite differences in the magnitude of the adjustment being slightly larger, the same holds for the reaction of import prices to a one unit shock in the exchange rate (panel a). Most importantly, both specification provide evidence for an incomplete but substantial and statistically significant adjustment of import prices and a weak reaction of consumer prices.

7 Changes in the degree of pass-through

The sample period considered is particularly interesting for the analysis of the ERPT as it contains two exogenous regulatory shocks directly affecting the *EURCHF* exchange rate. The shocks were both caused by the SNB and represent drastic policy changes with regards to the foreign exchange rate market. The first regulatory shock consists of the introduction of a *EURCHF* minimum exchange rate of 1.20 Swiss Francs per Euro (hereafter floor) in September 2011. This unconventional policy was implemented as a response to appreciation pressures on the Swiss currency, both in real and nominal terms. The Swiss Franc had already consistently appreciated over decades. The appreciation was intensified even more by both the financial crisis of 2008 and the Euro crisis in summer 2011, which both caused safe heaven flows into the Swiss currency to increase substantially. At the time of the floor introduction, the SNB stated that "the current massive overvaluation of the Swiss franc poses an acute threat to the Swiss economy and carries the risk of a deflationary development." and that, therefore, it "will enforce this minimum rate with the utmost determination and is prepared to buy foreign currency in unlimited quantities." (Swiss National Bank, 2011, p.1).

The period of exchange rate stability stemming from the floor ended abruptly with the second intervention, the discontinuation of the floor in January 2015. The timing of the decision of the SNB to suddenly discontinue the floor was motivated by changing conditions of international financial markets. In particular, the increasing differences in monetary policy actions between the ECB and the Federal Reserve prompted the decision. In connection to the preceding policy hikes in the United States, the SNB press release of the discontinuation announcement in January 2015 stated that "recently, divergences between the monetary policies of the major currency areas have increased significantly" and it was concluded that "enforcing and maintaining the minimum exchange rate for the Swiss Franc against the Euro is no longer justified." (Swiss National Bank, 2015, p.1).

The two interventions are noticeable when visually inspecting the exchange rate time series, both in levels and in log-first differences (see Figure 7.1 and panel (a) of Figure 7.2). There are many reasons to believe that the interventions affected the ERPT mechanism to Swiss import and consumer prices. First of all, as mentioned in Section 2, exchange rate volatility is thought to be negatively correlated with the intensity of the pass-through. The higher the exchange rate volatility, the more should firms prefer to continuously adjust their markups rather than their prices (Mann, 1986). This correlation has also been widely confirmed empirically via cross country studies (see, for instance Cheikh & Louhichi,

Figure 7.1: Introduction and Discontinuation of the *EURCHF* Minimum Exchange Rate

Notes: The figure shows the *EURCHF* nominal exchange rate in levels over the sample period 2000:01 to 2019:09 (black solid line). Vertical red dashed lines represent SNB's interventions: the first represents the introduction of the 1.20 Swiss Francs per Euro floor (September 2011), the second the discontinuation of the same floor (January 2015). Source: SNB.

2015; Corsetti et al., 2008; McCarthy, 2007). The two interventions directly affected the *EURCHF* exchange rate volatility, as they represent transitions from a floating exchange rate regime to an almost fixed exchange rate one and vice versa. Moreover, ERPT is similarly affected by the persistance of exchange rate shocks. That is, if firms believe a shock to be permanent, they are more likely to pass it to the consumers by changing prices (see An and Wang (2012), but also Jašová et al. (2019)). Given the commitment of the SNB to maintain the floor, the first shock is likely to have been perceived as a permanent one. The same probably holds for the second shock as well: the SNB clearly stated how a continuation of the floor had become unjustified, if not unstustainable. It is therefore likely that firms started to expect the sudden appreciation of the Swiss Franc caused by the discontinuation of the floor to be quite persistent, if not even increasing (as during pre floor times).

This section aims to investigate the extent to which these two interventions affected the ERPT to import and consumer prices in Switzerland. To this end, in a first step, a series of Chow tests of time invariance of different groups of parameters are estimated in order to assess if and how the two interventions impacted the pass-through dynamics, as described by a VAR model. Based on tests results, in a second step, an intervention model is specified and used to compare pass-through dynamics across the three different exchange rate regimes, namely pre floor regime (from 2000:01 to 2011:08), floor regime (from 2011:09 to 2014:12) and post floor regime (from 2015:01 to 2019:09). First, however, the three time series of most relevance for pass-trough dynamics, namely the exchange rate, the IPI and the CPI, are analysed separately with respect to the two interventions by means of Bai and Perron tests (see Bai & Perron, 1998).

7.1 Analysis of single time series

Despite the study aiming to analyse the effect of the two interventions on Swiss ERPT dynamics, i.e. on the entire system considered, an investigation of potential effects of the interventions on the single time series can be of interest and of help for better understanding their overall effects. Therefore, this section analyses the three for ERPT dynamics most relevant time series, namely the exchange rate, the Swiss IPI and the Swiss CPI, separately with regards to the introduction of the floor and its subsequent discontinuation by means of Bai and Perron tests.

Bai and Perron (1998) proposed a testing procedure for the detection of multiple structural breaks at unknown dates in linear regression models. A nice feature of their procedure is that it allows to estimate not only the number of breaks but also the dates of the breaks. Moreover, it is able to test for pure structural breaks, i.e. breaks in all regression's parameters, as well as partial structural breaks, i.e. breaks in the constant of the regression only. In both cases, their model allows for heterogeneity in the error terms. However, it does not provide a method to estimate it (pp.49-51).

In essence, their method consists of estimating the unknown coefficients of the linear regression together with the break points by means of the ordinary least-square principle under fairly general conditions of the DGP and of the error terms. Once the estimations are completed, inference on the number and position of breaks is based on certain information criteria[1] and the statistic values of three types of tests. The first tests no breaks against a fixed predefined number of breaks (p.57). The second type, a so called double maximum test, instead of requiring a specific number of breaks under the alternative hypothesis, tests no breaks against an unknown number of breaks (other than zero) given some upper bound on the potential number of breaks (pp.58-59). Finally, the third type consists of a sequential test, which tests l breaks against the alternative hypothesis of $l + 1$ breaks. Starting with testing zero breaks against one, the number of breaks chosen is the one implied by the first null hypothesis that cannot be rejected (pp.64-65). All three types of tests have non-standard asympotic distributions for which Bai and Perron (1998) provide critical values computed through simulations.

As suggested by the authors, when selecting the number of breaks based on the above mentioned tests, the sequential test is to be considered first. If the null hypothesis of zero breaks is rejected, one should continue with the sequential testing procedure and select the number of breaks accordingly. If the null is not rejected, one should look at the double

[1] Such as the BIC or the modified Schwarz criterion (LWZ).

maximum test. If the hypothesis of zero breaks cannot be rejected there as well, then the procedure stops and it should be concluded that it found no breaks. If, however, the double maximum test is able to reject the hypothesis of zero breaks, authors suggest to go back to the sequential testing procedure but, this time, starting from one break against two (Bai & Perron, 2003, pp.15-16).

Here, the idea is to exploit the described procedure to test the three mentioned series for multiple structural breaks (both pure and partial) separately and as they appear in the baseline VAR model, i.e. in their logarithm and in first differences, so to check if and how the two interventions impacted them singularly. To do that, similarly as for the ADF tests implemented in Section 3.2, the series are modelled as AR(p) processes, with number of lags selected by means of AIC and BIC in combination with well behaved residuals. More details on the Bai and Perron (1998) testing methodology and on its implementation for the case considered here can be found in Appendix E.

Figure 7.2 depicts the three series of interest in log first differences. For both the exchange rate and the IPI, the two interventions seem to coincide with a spike. For the CPI, on the other hand, no clear spike is visible at any intervention date. Furthermore, for both IPI and CPI, some spikes appear around the financial crisis of 2008.

In contrast with expectations, the testing procedure is not able to reject the null hypothesis of no structural breaks for the exchange rate series. Both sequential testing and double maximum test suggest this when looking at pure and partial structural breaks. More in line with expectations are results for the IPI and the CPI series. While rejecting the presence of any pure structural break, Bai and Perron tests for the CPI series identify one partial structural break in 2008:08[2], most likely caused by the financial crisis. For the IPI series, while no partial breaks are identified, two pure breaks are: the first in 2008:08 (again, most likely representing the financial crisis), the second one in 2015:04. The latter is likely to corresponds to the discontinuation of the floor in 2015:01[3] Details on test and critical values can be found in Appendix E.

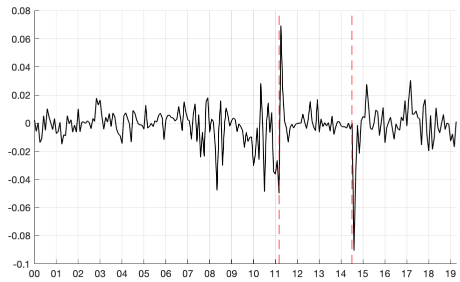
All in all, results provide no or little evidence for the two interventions representing a structural break in the single series considered. Moreover, they seem to indicate that the financial crisis has caused a structural break in the Δcpi_t and Δipi_t series. With that said, results are to be interpreted with caution as based on many assumptions. For instance, if the underlying AR(p) processes are misspecified, test results might be misleading. Nevertheless, test results with regards to the financial crisis should be taken seriously. If the crisis represents a structural break in two of the, for the ERPT, most relevant series, it

[2] Results based on the sequential testing procedure. BIC suggests the presence of one break too, while the LWZ suggests no breaks.

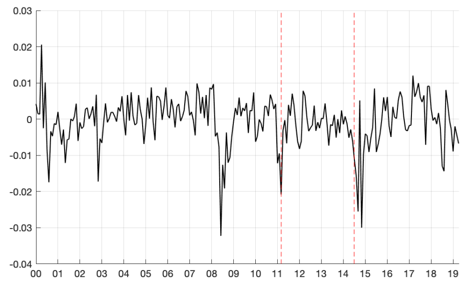
[3] Especially considering the fact that the break date represents an estimate with confidence bounds given by a range of dates around the estimated one.

might also represent one for overall ERPT dynamics, in which case it could compromise estimation results if not adequately addressed. In order to better understand the role of the financial crisis for the ERPT dynamics, the baseline specification has been estimated with a dummy variable for the financial crisis[4] as additional exogenous regressor. Accumulated orthogonalised impulse responses for model specification with and without financial crisis dummy variable are depicted in Figure F.1 in Appendix F. As one can see, estimates are almost identical, suggesting that the financial crisis seems not to affect ERPT estimates to an important extent.

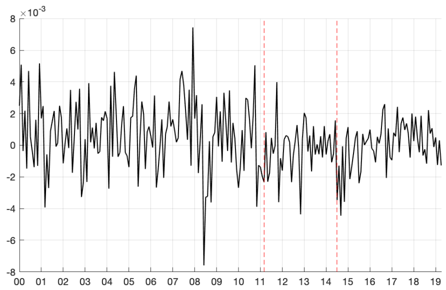
Figure 7.2: Visual Inspection of Series in Log-First Differences



(a) *EURCHF* Exchange Rate



(b) Import Price Index



(c) Consumer Price Index

Notes: The Figure depicts the *EURCHF* nominal exchange rate (panel a), the Swiss IPI (panel b) and the Swiss CPI (panel c) (black solid lines), all in log first differences over the sample period 2000:01 to 2019:09. Swiss IPI and Swiss CPI have been seasonally adjusted by means of the Census-X12 procedure. Vertical red dashed lines represent SNB's interventions: the first represents the introduction of the 1.20 Swiss Francs per Euro floor (September 2011), the second the discontinuation of the same floor (January 2015). Source: Author's rendering of SNB and SFSO data (2020).

After having analysed the impact of the interventions on the single series separately,

[4] The dummy variable has been set to zero except for September 2008, where it has been set to one. September 2008 corresponds to the failure of the Lehman Brothers Bank and is therefore considered the starting date of the global financial crisis.

in what follows, their potential effects on the overall ERPT dynamics governed by the entire system will be assessed. Indeed, the evidence provided by the Bai and Perron tests does not necessarily exclude the possibility that the two interventions impacted the relationship between the three series analysed. Moreover, tests were not able to identify structural breaks in the variance of the processes. These are therefore as well not to be excluded. In light of this, an analysis of potential effects on the entire system is more than justified.

7.2 The intervention model and Chow tests analysis

Given the nature of the two interventions and theoretical considerations on factors affecting ERPT, it seems reasonable to expect the interventions to have impacted both the volatility and persistence of changes in the exchange rate as well as their transmission to import and consumer prices. In other words, it is expected that the interventions affected both the volatility of the shocks of the system (especially those to the exchange rate), i.e. the variance covariance matrix Σ_u, as well as the IRF of import and consumer prices, i.e. the coefficient matrices A_i of the reduced form. In this context, a so called intervention model together with a series of Chow tests represent a potential approach to empirically test these expectations.

As explained by Lütkepohl (2005), intervention models describe the case where a particular stationary DGP is in operation until a certain period, while an another, different DGP operates afterwards (p.604). Here, it is assumed that three potentially different DGPs exist, one for each regime. Then, the corresponding intervention model is constructed by defining three "intervention dummy variables" in accordance with the regimes. Formally, these dummy variables are defined as:

$$n_{1t} = \begin{cases} 1 & \text{if } t \leq 2011{:}08 \\ 0 & \text{otherwise} \end{cases}$$

$$n_{2t} = \begin{cases} 1 & \text{if } 2011{:}09 \leq t \leq 2014{:}12 \\ 0 & \text{otherwise} \end{cases}$$

$$n_{3t} = \begin{cases} 1 & \text{if } t \geq 2015{:}01 \\ 0 & \text{otherwise} \end{cases}$$

Then, the intervention model can be such that it implies either a simple discrete change in the mean or a change in the intercept. In the first case, the system is assumed to experience only a one time jump in its mean at the regime switching period. In contrast, in the second case, the system is supposed to react slowly to the intervention, i.e. over more than one period. A change in the intercept is plausible when a smooth adjustment of the general economic conditions seems more realistic then a discrete change, so that the system should be allowed to gradually approach a new equilibrium (Lütkepohl, 2005, pp.605-609). Given the type of interventions analysed here and the fact that they implied transitions from a floating to an almost fixed exchange rate regime and vice versa, a model implying an intervention in the intercept seems more appropriate. Indeed, Bonadio et al. (2020), in their study on currency of invoicing and ERPT in Switzerland, state in connection to the discontinuation of the floor that "the appreciation occurred in a stable macroeconomic environment and that the Swiss economy quickly settled to a new equilibrium after the shock."(p.6). The VAR(p) intervention model to be specified is then the following[5]:

$$y_t = \nu_t + A_{1t}y_{t-1} + ... + A_{pt}y_{t-p} + u_t \tag{7.1}$$

where $u_t \sim (0, \Sigma_{ut})$. Also, define

$$\begin{aligned} B_t &= [\nu_t, A_{1t}, ..., A_{pt}] \\ &= n_{1t}[\nu_1, A_{11}, ..., A_{p1}] + n_{2t}[\nu_2, A_{12}, ..., A_{p2}] + n_{3t}[\nu_3, A_{13}, ..., A_{p3}] \\ &= n_{1t}B_1 + n_{2t}B_2 + n_{3t}B_3 \end{aligned}$$

with dimensions $[K \times (Kp + 1)]$, and

$$\Sigma_{ut} = E(u_t u_t') = n_{1t}\Sigma_{u1} + n_{2t}\Sigma_{u2} + n_{3t}\Sigma_{u3}$$

[5] Again, for the purpose of simplification, exogenous terms have been ignored in derivations.

with dimensions $K \times K$. The great feature about this specification is that it allows for all parameters to change across the three regimes, without necessarily implying that. In other words, the stationary time invariant VAR model of the baseline specification is nested in model 7.1 (Lütkepohl, 2005, p.586).

The hypothesis of the two interventions having caused both the variance covariance of the error terms as well as the effect of shocks on the system to change can be tested by means of Chow tests on model 7.1. These are based on obtaining estimates under various types of restrictions on different sets of parameters, thereby testing the time invariance of different groups of parameters. Chow tests are usually implemented as LR tests, of which the general form is

$$\lambda_{LR} = 2[\log l(\tilde{\delta}) - \log l(\tilde{\delta}_r)]$$

where $\tilde{\delta}$ represents the unconstrained ML estimator and $\tilde{\delta}_r$ the restricted one, both obtained by maximizing the likelihood function under the null hypothesis. If the null hypothesis holds and under general conditions, λ_{LR} has a χ^2 distribution with degrees of freedom equal to the number of linearly independent restrictions imposed. Indeed, it can be shown that under certain restrictions, ML estimation of model 7.1 becomes fairly trivial (Lütkepohl, 2005, pp.595-596).

Here, the idea is to test three different null hypotheses implying the time invariance of different sets of parameters against the single alternative hypothesis of all parameters changing across regimes. Then, a rejection of all null hypotheses would be taken as suggestive evidence in favour of the two interventions having changed both variance of the shocks as well as their effect on prices. Following Lütkepohl (2005), the three null hypotheses tested are formally summarised as (pp.595-601):

H_0^1 : stationary model with constant parameters. Formally:

$$H_0^1\text{: } B_i = B_1 \text{ and } \Sigma_{ui} = \Sigma_{u1} \text{ for } i = 2, 3$$

H_0^2 : time varying[6] intercept, constant coefficients and variance. Formally:

$$H_0^2\text{: } \nu_t = \textstyle\sum_{i=1}^{3} n_{it}\nu_i;\ A_i = A_1 \text{ and } \Sigma_{ui} = \Sigma_{u1} \text{ for } i = 2, 3$$

H_0^3: time varying intercept and coefficients, constant variance. Formally:

$$H_0^3\text{: } B_t = [\nu_t, A_1] = \textstyle\sum_{i=1}^{3} n_{it}B_i \text{ and } \Sigma_{ui} = \Sigma_{u1} \text{ for } i = 2, 3$$

H_0^1 implies nothing else than the baseline specification of Section 5. Under H_0^2, the interventions caused only the mean of the system to change, while under H_0^3 they changed

[6] Here, time varying refers to changing across regimes.

both the mean of the system and the transmission mechanisms of the shocks to prices but not the variance of the shocks.

The single alternative hypothesis of all parameters changing across regimes is formally expressed as:

H^a : time varying intercept, coefficients and variance. Formally:

$$H^a\text{: } B_t = [\nu_t, A_t] = \sum_{i=1}^{3} n_{it} B_i \text{ and } \Sigma_{ut} = \sum_{i=1}^{3} n_{it} \Sigma_{ui}$$

Parameters under the different hypotheses are estimated by means of ML and then used to compute the statistic values of the LR tests. Candelon and Lütkepohl (2001) showed how, in small samples, the distribution of Chow test statistics may differ substantially from their theoretical asymptotic χ^2 distributions[7]. Therefore, is is strongly suggested to implement bootstrap versions of the tests as well. Following the advice, for each test, a bootstrap version is computed alongside the theoretical one. To conserve space, technical aspects of the ML estimations and of the computation of LR test values as well as their bootstrap versions are reported in Appendix G. Here, results only are shown.

Table 7.1: Chow Tests Results

H_0	Test Statistic	Degrees of Freedom (DF)	$\chi^2(0.99, DF)$	Monte Carlo p-value
H_1^0	395.60	190	238.27	0.0018
H_2^0	472.65	180	227.06	0.0002
H_3^0	224.06	30	50.89	0.0002

Notes: The table depicts Chow tests results. All null hypotheses are tested against the same alternative hypothesis of all parameters changing across regimes. Critical values for asymptotic results of the test are at the 1% significance level. Parametric bootstrap p-values are computed based on 5000 Monte Carlo simulations.

Table 7.1 depicts test results. The three null hypotheses are clearly rejected. Bootstrap results confirm this. Therefore, the model implied by the alternative hypothesis seems to be the preferred one. With that said, it is important to note that not all potential different sets of parameters have been tested for time invariance against the alternative of all parameters changing across regimes. For instance and in particular, the null hypothesis implying a time varying variance covariance structure but constant coefficients and intercept has not been tested against the alternative. The reason behind this is mainly that, as explained by Lütkepohl (2005), estimation procedures for this type of model are quite cumbersome due to the presence of nonlinearities. Consequently, the computation of bootstrap test values becomes quite challenging as well, since based on simulating the

[7] The small sample concern is even stronger in the context of intervention models as the information about the parameters is often limited within the time frame prior to the intervention, making it difficult to argue that it increases as the sample size goes to infinity (Lütkepohl, 2005, p.606).

DGP under the null hypothesis. Moreover, from an economic perspective, VAR models with extraneously specified volatility changes don't seem to be well suited for the case considered here. As explained by Kilian and Lütkepohl (2017), heteroskedasticity in VAR models is often included to solve the identification problem through a purely statistical procedure, i.e. without the help of economic intuition. Indeed, Rigobon (2003) proposed this method as a way to solve the identification problem when all other methods appear unreasonable or not supported by economic reasoning. However, the restrictions by the unconditional heteroskedasticity only suffice for identification if the coefficients of the contemporaneous effects are assumed to be fixed across regimes (Kilian & Lütkepohl, 2017, p.525). This assumption being quite restrictive, Bacchiocchi and Fanelli (2015) consider the possibility of time varying contemporaneous effects by implementing a "new" identification procedure, in which identification is still based on heteroskedasticity but at the same time allows for changes in the other coefficients. The pitfall of their approach is that when effects are allowed to change, the restrictions imposed by heteroskedasticity no longer suffice for identification. That is, additional restrictions regarding how and which coefficients are allowed to change across regimes are to be imposed. This can be problematic as, first, such assumptions might not be easy to justify from an economic perspective and, secondly, if there are multiple regimes, the number of additional restrictions to be imposed becomes extremely large. For these reasons, even models such as those proposed by Bacchiocchi and Fanelli (2015) are still considered quite restrictive (Kilian & Lütkepohl, 2017, p.525). Moreover, they represent situations in which a change in the contemporaneous effects results in a more general change of the IRF despite the reduced form parameters not being regime dependent, which reflects a rather unrealistic scenario in the majority of cases.

Due to such considerations, test results and economic intuition being quite supportive for a model with all parameters changing across regimes, an intervention model in line with H^a has been estimated. The results of such model are discussed in the next section.

7.3 Results of the intervention model

To adress potential breaks caused by the two SNB interventions, an intervention model of the form 7.1 with all parameters changing across regimes has been estimated by ML. For the order selection of the model, Lütkepohl (2005)'s approach has been followed once more: as 3 lags seem to have performed well for the entire sample period in the baseline specification, the same number of lags is included in the intervention model.

7.3.1 Exchange rate pass-through to import prices

Figure 7.3 depicts accumulated IRF of import prices to a one percent depreciation of the Swiss Franc against the Euro for the three regimes separately. Table 7.2 reports the corresponding estimates in numbers. For the purpose of comparison, the table reports also the estimates of the baseline specification. Estimates turn out to be significantly different across regimes, suggesting that the interventions had some effects on the transmission mechanism of exchange rate changes to import prices.

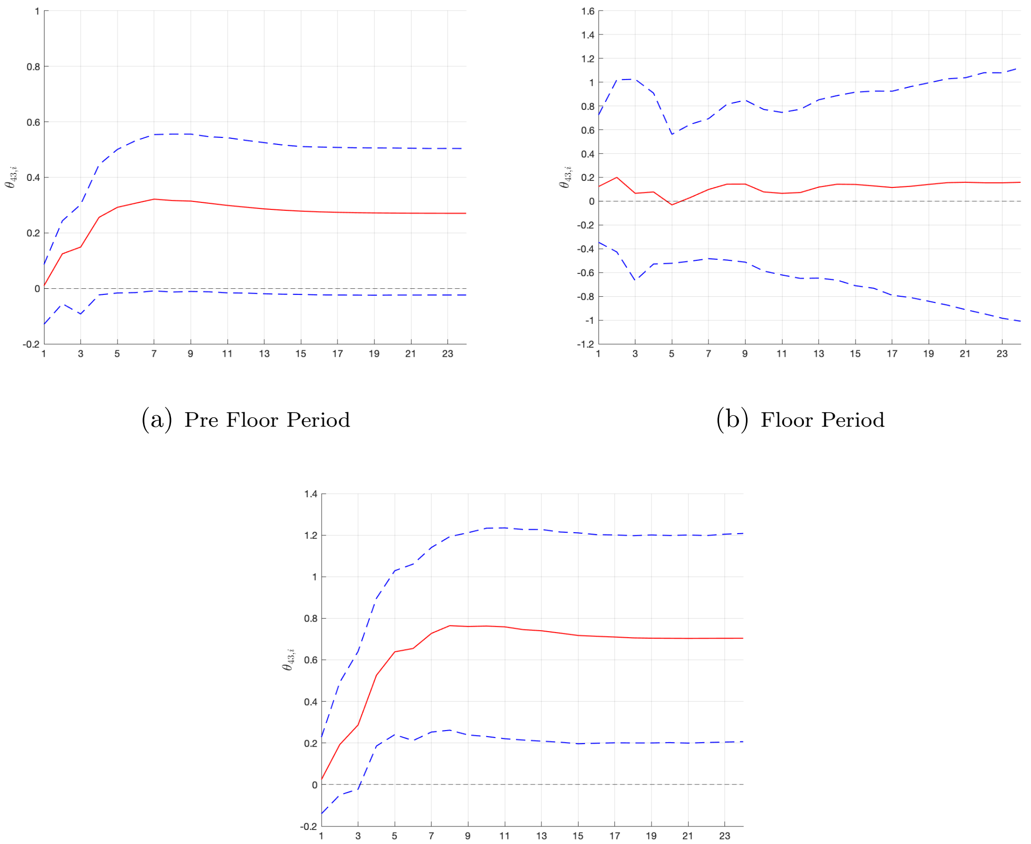
Table 7.2: Exchange Rate Pass-Through to Import Prices

	Pre floor period		**Floor period**		**Post floor period**		Baseline estimates	
Period i	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	IRF	Acc. IRF
1	0.009790	0.009790	0.122759	0.122759	0.024311	0.024311	0.039279	0.039279
3	0.024504	0.149071	-0.133953	0.065940	0.094374	0.286900	0.038015	0.223463
6	0.015023	0.307188	0.061620	0.030454	0.016149	0.654815	0.020968	0.468933
9	-0.001985	0.314529	0.000688	0.143642	-0.004031	0.760492	-0.002651	0.497144
12	-0.006360	0.292603	0.007109	0.072725	-0.013369	0.745411	-0.010349	0.467162
15	-0.003529	0.278491	-0.001911	0.140567	-0.011614	0.717277	-0.007554	0.439763
18	-0.001310	0.272793	0.010223	0.124971	-0.004244	0.705550	-0.003647	0.424909
21	-0.000388	0.270968	0.003313	0.158965	-0.000479	0.703088	-0.001211	0.419203
24	-0.000109	0.270440	0.004380	0.158835	0.000172	0.703830	-0.000168	0.417899

Notes: The table shows orthogonalised impulse responses and accumulated orthogonalised impulse responses of Δipi_t to a one unit shock in Δe_t. Pre floor, floor and post floor period correspond to 2000:01 to 2011:08, 2011:09 to 2014:12 and 2015:01 to 2019:09 respectively. The first six columns show estimates based on the intervention model 7.1 estimated with 3 lags for both endogenous and exogenous variables, the last two columns show estimates based on the baseline VARX(3,3) model.

When looking at the floor period (panel b), the shape of the accumulated IRF is rather off its usual humped one and they appear not to be statistically significant (as suggested by confidence bounds including the zero). Considering the extremely low volatility of the exchange rate during this period, these results are rather unexpected. However, the deflationary period Switzerland was going through before the introduction of the floor could be a potential explanation. Indeed, when estimating monthly ERPT elasticities to import prices for Switzerland, Fleer et al. (2016) find that the ERPT to import prices experienced a period of increase at the beginning of 2011. They suggest deflationary processes as a response to the Swiss Franc appreciation as one potential explanation for such results (p.21). After the introduction of the floor, their monthly estimates start decreasing. Bringing this in connection to panel b of Figure 7.3, a potential explanation could be that once the floor was introduced, downward pressure on prices was alleviated by the minimum exchange rate, leading to a situation in which prices could remain rather constant. Furthermore, Gagnon and Ihrig (2004) show how a strong commitment of the monetary policy authority towards stabilising inflation is associated with lower ERPT. In this sense, results point towards the intervention having been effective in reducing or stopping deflation. On the other hand, it might simply be that, due to the floor, the variance in the exchange rate series is too limited to provide meaningful estimates.

Figure 7.3: Exchange Rate Pass-Through to Import Prices



(a) Pre Floor Period

(b) Floor Period

(c) Post Floor Period

Notes: The figure shows accumulated orthogonalised impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a one-unit shock in the exchange rate based on model 7.1 and 3 lags for both endogenous and exogenous variables. Blue dashed lines represent 95% standard residual-based recursive design bootstrap confidence bounds.

In contrast, pre and post floor periods (panel a and c) provide estimates in line with expectations. That is, ERPT to import prices is humped shaped. Compared to baseline estimates, the pass-through in the pre floor period is substantially lower and no longer statistically different from zero[8] while, in the post floor period, substantially higher and statistically different from zero. The maximum is still reached after nine months approximately, but amounts to 0.31 in the pre floor period and to 0.76 in the post floor period compared to 0.50 estimated with the baseline model (see Table 7.2). These results could be explained by the fact that when using the intervention model, the effect of the appreciation shock caused by the discontinuation of the floor is accounted for only by the estimates of the third regime (panel c), where the ERPT to import prices results to be almost complete, reaching a level of 70 percent after two years. Indeed, as explained by Jašová et al. (2019), large exchange rate movements create larger reactions in prices

[8] With thatt said, 90% confidence bounds are still different from zero.

as they are able to overcome the menu costs that often block firms from adjusting their prices. It is therefore likely that the Swiss ERPT to import prices during "normal" times, i.e. such as those prior to the introduction of the floor, is lower than the one estimated when considering the entire sample and that, since the discontinuation of the floor, the ERPT to import prices has significantly increased.

7.3.2 Pass-through of import prices to consumer prices

Table 7.3: Pass-Through of Import Prices to Consumer Prices

	Pre floor period		**Floor period**		**Post floor period**		Baseline estimates	
Period i	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	IRF	Acc. IRF
1	0.203039	0.203039	0.286855	0.286855	0.104285	0.104285	0.180346	0.180346
3	0.041599	0.280431	0.034872	0.325789	-0.018532	0.136188	0.026027	0.247098
6	0.002573	0.284447	0.029349	0.125708	-0.001688	0.129923	0.004262	0.271149
9	-0.005479	0.271315	-0.030109	0.238880	-0.002158	0.118152	-0.002379	0.265205
12	-0.003229	0.258694	-0.000495	0.172297	-0.001355	0.111472	-0.002860	0.255442
15	-0.001178	0.253697	0.004608	0.224178	0.000258	0.111337	-0.001596	0.249313
18	-0.000336	0.252058	0.002666	0.194873	0.000289	0.112286	-0.000626	0.246595
21	-0.000089	0.251610	-0.001255	0.213508	0.000186	0.112789	-0.000147	0.245777
24	-0.000027	0.251483	0.000492	0.204339	0.000018	0.112976	0.000017	0.245712

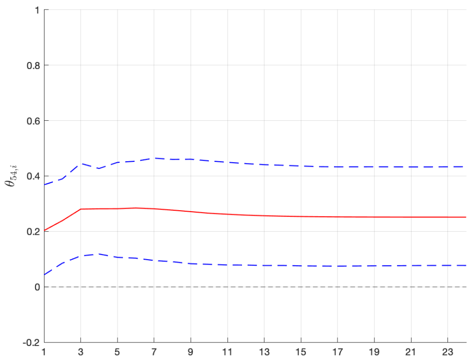
Notes: The table shows orthogonalised impulse responses and accumulated orthogonalised impulse responses of Δcpi_t to a one unit shock in Δipi_t. Pre floor, floor and post floor period correspond to 2000:01 to 2011:08, 2011:09 to 2014:12 and 2015:01 to 2019:09 respectively. The first six columns show estimates based on the intervention model 7.1 estimated with 3 lags for both endogenous and exogenous variables, the last two columns show estimates based on the baseline VARX(3,3) model.

Similarly as for the ERPT to import prices, the pass-through of import prices to consumer prices differs across the three regimes and appears not to be significantly different from zero when the floor period is considered (see Figure 7.4 and Table 7.3). Surprisingly in contrast with the ERPT to import prices, however, is the comparison of pre and post floor periods with baseline estimates. Here, pre floor period estimates (panel a) are higher than baseline estimates, while post floor period estimates (panel c) are smaller. In fact, pre floor accumulated IRF are close or even over the average share of imported goods in the CPI of 25 percent, while post floor accumulated IRF are well below. This could reflect a decrease in the share of imported goods in the CPI basket, a change in the amount of intermediate imported goods or a change in the response of prices of domestic goods to changes of prices of imported ones. A decrease in the share of imported goods sounds rather unlikely given the fact that the domestic currency appreciated. Indeed, comparing the share of imported goods across the years shows that it stayed relatively constant around approximately 25 percent of the CPI[9]. In light of this, the difference between pre and post floor period is more likely to be driven by direct effects on domestic goods.

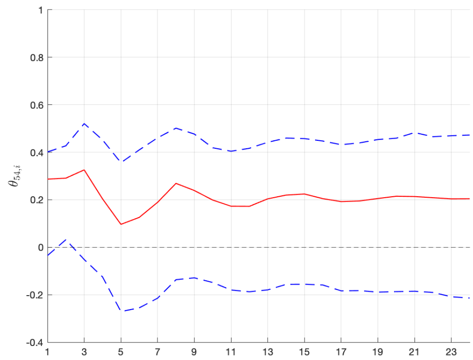
[9] For instance, in 2014, before the discontinuation of the floor, imported goods amounted to 26.74% of the CPI, at the end of 2016, approximately one year after the discontinuation, they amounted to 24.45% and at the end of 2019 to 25.3%. Source: SFSO's weighting of the CPI (Landesindex der Konsumentenpreise - Gewichtung) for the years from 2014, 2016 and 2019.

One would expect domestic producers to decrease prices of domestic goods in response to imported goods becoming cheaper in relative terms due to the appreciation of the Swiss Franc. Results, however, suggest the contrary, i.e. that domestic producers most likely did not adjust their prices despite import goods becoming relatively less expensive.

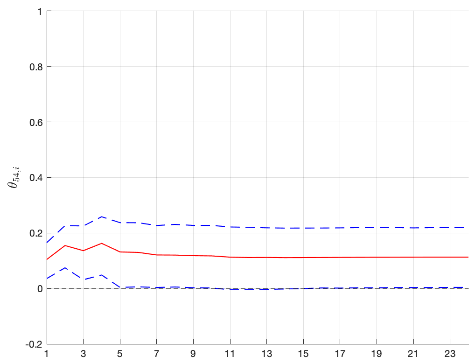
Figure 7.4: Pass-Through of Import Prices to Consumer Prices



(a) Pre Floor Period



(b) Floor Period



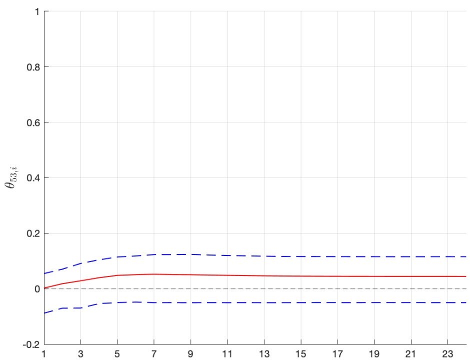
(c) Post Floor Period

Notes: The figure shows accumulated orthogonalised impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a one-unit shock in import prices based on model 7.1 and 3 lags for both endogenous and exogenous variables. Blue dashed lines represent 95% standard residual-based recursive design bootstrap confidence bounds.

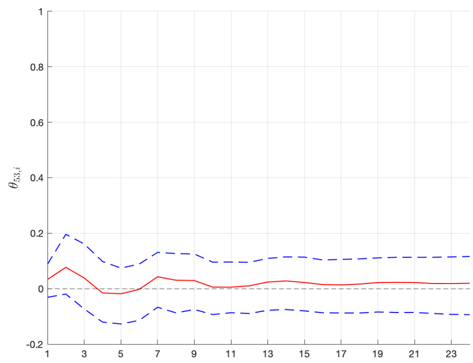
7.3.3 Exchange rate pass-through to consumer prices

Figure 7.5 graphically shows accumulated IRF of a one percent shock in the exchange rate to consumer prices. Table 7.4 reports the same estimates in numbers. In contrast to the other two pass-troughs, the ERPT to consumer prices does not substantially differ across regimes and, most importantly, does not appear to be significantly different from zero for any of the three periods. Considering the already extremely low baseline estimates, this is not too surprising. Nevertheless, given the almost complete ERPT to import prices estimated separately for the post floor period, such low and insignificant estimates might be surprising for the third regime only (panel c). On the other hand, these might be explained by the extremely low (lower than baseline estimates and pre floor period estimates) pass-through of import to consumer prices estimated for the same period.

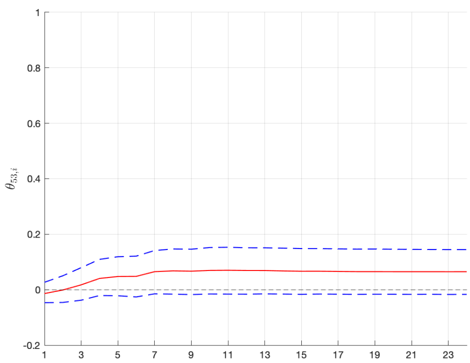
Figure 7.5: Exchange Rate Pass-Through to Consumer Prices



(a) Pre Floor Period



(b) Floor Period



(c) Post Floor Period

Notes: The figure shows accumulated orthogonalised impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a one-unit shock in the exchange rate based on model 7.1 and 3 lags for both endogenous and exogenous variables. Blue dashed lines represent 95% standard residual-based recursive design bootstrap confidence bounds.

Overall, results suggest that both the introduction as well as the discontinuation of the floor have impacted ERPT dynamics to import and consumer prices, as estimates differ substantially across the three regimes. Other interesting insights are revealed. First, results point towards the introduction of the floor having been effective in stopping the deflationary development caused by the previous substantial and continuous appreciation of the Swiss Franc. Secondly, it emerges that while the exchange rate pass-through to import prices is significantly higher in the post floor period, pass-through of import prices to consumer prices is higher in the pre floor period. This most likely explains why the ERPT to consumer prices in the post floor period appears lower than in the pre floor

Table 7.4: Exchange Rate Pass-Through to Consumer Prices

	Pre floor period		**Floor period**		**Post floor period**		Baseline estimates	
Period i	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	**IRF**	**Acc. IRF**	IRF	Acc. IRF
1	0.002355	0.002355	0.033262	0.033262	-0.013832	-0.013832	0.006852	0.006852
3	0.010458	0.028539	-0.038463	0.038625	0.018650	0.017531	0.012311	0.041553
6	0.002454	0.050514	0.015896	-0.002263	0.000464	0.048055	0.003040	0.067291
9	-0.000430	0.050298	-0.001246	0.029282	-0.001084	0.066829	-0.000312	0.073161
12	-0.000952	0.047245	0.004935	0.010453	-0.000745	0.069313	-0.001496	0.069108
15	-0.000464	0.045312	-0.005676	0.022293	-0.001087	0.066577	-0.001069	0.065310
18	-0.000168	0.044572	0.002868	0.016695	-0.000653	0.065248	-0.000529	0.063173
21	-0.000047	0.044344	-0.000620	0.022106	-0.000126	0.064919	-0.000180	0.062332
24	-0.000013	0.044280	0.001056	0.019597	0.000017	0.064916	-0.000028	0.062129

Notes: The table shows orthogonalised impulse responses and accumulated orthogonalised impulse responses of Δcpi_t to a one unit shock in Δe_t. Pre floor, floor and post floor period correspond to 2000:01 to 2011:08, 2011:09 to 2014:12 and 2015:01 to 2019:09 respectively. The first six columns show estimates based on the intervention model 7.1 estimated with 3 lags for both endogenous and exogenous variables, the last two columns show estimates based on the baseline VARX(3,3) model.

period and not statistically significant, despite the ERPT to import prices being close to complete. This seems to suggest that, in contrast to the conclusions based on baseline results, for the post floor period, the low pass-through to consumer prices seems to be caused by a low pass-through of import prices to consumer prices rather than sticky import prices. In fact, the latter appear to adjust substantially to changes in the exchange rate.

8 Discussion

After having presented all empirical results, the following section aims to discuss them both in terms of potential policy implications as well as limitations.

8.1 Policy implications

In the context of flexible exchange rates, a substantial ERPT allows the PPP to hold to a certain extent and, thus, to protect consumers from excessively paying a product in the case of a domestic currency appreciation. On the other hand, as explained by Stulz (2007), a low ERPT provides greater freedom for pursuing an independent monetary policy and makes it easier to control inflation (p.23). Indeed, policymakers must be able to prevent changes in relative prices stemming from exchange rate movements and fuelling a continuous inflationary or deflationary process. For the Swiss case specifically, as already mentioned more than once, due to the safe heaven nature of the Swiss Franc, the risk of deflation is high and its mitigation has been on top of the agenda of the SNB for nearly the past ten years. The intensity of the pass-through both in the short- and long-run is therefore important to understand the impact of exchange rate movements not only on prices but also on quantities and, hence, welfare (Burstein & Gopinath, 2013, p.392). Moreover, independently of its degree, awareness about the pass-through level is of crucial importance for inflation forecasts, upon which monetary policy decisions are based.

Based on the empirical results of this paper, policy implications to be derived are the following. Monetary policy authorities should be aware of the fact that, despite consumer prices not reacting much to changes in the exchange rate, import prices do adjust substantially to such changes, especially since the beginning of 2015. This might make the ERPT to consumer prices not a decisive factor for short-term inflation forecasts, but makes the monitoring of the relationship between import and consumer prices of high relevance, since, should this relationship change, exchange rate changes could substantially be passed to consumer prices as well, potentially leading to a situation of deflationary risk similar to as before the floor introduction. In connection to this, a further understanding of the business dynamics of Swiss importers might be of special interest, as well as of the degree of substitutability between imported goods and domestic goods and the pricing strategy of domestic producers. For the same reasons as for the importance of monitoring the relationship between import and consumer prices, it would also be crucial to assess whether the increase in the pass-through to import prices following the discontinuation of the floor is temporary or whether it represents a new equilibrium level. Such an assessment would be relevant also in light of the fact that a future reintroduction of a floor might be taken into consideration (in Switzerland, but potentially also in other small open economies), in which case the effects of a subsequent interruption of such floor should be taken into account.

8.2 Limitations

Despite the application of VAR models being popular in the field of Macroeconomics, several limitations to the underlying results should be addressed. First, there is the risk of having omitted a confounding variable, i.e. one that is correlated with both the response and the explanatory variables, in which case results would be inconsistent. This represents even more of a concern when it comes to specifying a VAR model in restricted samples, as one faces the tradeoff between including more variables and reducing the risk of omitted variable bias (OVB) at the cost of reducing the statistical degrees of freedom. Despite having included six important economic variables governing exchange rate dynamics, there is no way to test whether the OVB risk has been addressed completely. The lag order selection is subject to the same issue.

A further limitation regards the endogeneity problem of the variables included in the system. Despite the VAR and its underlying structural VAR models allowing for a certain degree of endogenous interaction between the variables, the identification of such interactions is based on identifying restrictions motivated by assumptions and economic reasoning. Therefore, results are sensitive to the choice of such restrictions. The many robustness checks with regards to the identification strategy, in particular the change of ordering and the computation of GIRF, suggest that endogeneity issues seem to not substantially undermine results. Said that, similarly as for the OVB risk, there is no test to check whether this issue has been addressed completely.

With regards to the intervention model specifically and the implementation of Chow tests, a limitation is represented by the fact that the tests implemented assume both interventions to have changed the same set of parameters. Given the nature of the interventions, this assumption seems rather plausible. However, there is no proper way to test it.

Finally, a limitation is to be found in the data used. More precisely, the output gap time series is the result of several estimation procedures. Not only do the monthly output gap values represent estimates rather than observations, but also the underlying monthly GDP values, as they had to be interpolated from the observed quarterly ones. Therefore, the series is likely to contain a lot of uncertainty and measurement error. With that said, given the purpose it was introduced for, rather than the actual value of the output gap, what is of most importance is the ability of the series to proxy for the progress of the economy (i.e. whether the economy is in a recession or in an expansion period). Therefore, to the extent that the series is able to accomplish this, the fact that it is the result of estimation procedures should not excessively undermine results.

9 Conclusions

This study aims to estimate the *EURCHF* exchange rate pass-through to Swiss import and consumer prices over the period from 2000:01 to 2019:09. With the introduction of a minimum *EURCHF* exchange rate in 2011 by the monetary policy authority and its subsequent discontinuation in 2015, during the period considered, the *EURCHF* exchange rate switched from freely floating to being capped and vice versa. In light of this, when estimating the exchange rate pass-through, the study also aims to address these two monetary policy interventions by investigating whether and to which extent they changed pass-through mechanisms.

To accomplish this, a VAR approach is implemented using monthly macroeconomic data. Identification is achieved by means of a standard Choleski decomposition. Through the imposed ordering of the variables, it is assumed that prices are set along the distribution chain, i.e. that the exchange rate and import prices have a contemporaneous effect on consumer prices, while consumer prices affect import prices and exchange rate only with a lag. Estimation of the size and speed of the pass-through is derived from impulse response functions of import and consumer prices to shocks in the exchange rate and import prices. The empirical analysis is carried out in two steps: first, a baseline model is estimated over the entire sample period, thereby ignoring the two interventions. In a second step, the effects of the two interventions are analysed by means of an intervention model allowing for parameters to change at the dates of the interventions. Based on a structural break analysis implemented through a series of Chow tests, both impulse responses and variance structure are allowed to change across exchange rate regimes.

The empirical investigation yields several interesting results. The evidence from the baseline analysis suggests that exchange rate changes are passed through quickly to import prices. In the long run, the related pass-through is incomplete but substantial and amounts to approximately 40 percent. Moreover, the transmission of import price shocks to consumer prices is surprisingly strong and close to the share of imported goods of approximately 25 percent. On the contrary, estimated exchange rate pass-through to consumer prices is weak. Taking these findings together, the conclusion to be drawn appears to be that the transmission of exchange rate changes to consumer prices seems to be primarily blocked by sticky import prices (i.e. by the incomplete exchange rate pass-through to import prices).

These results appeared robust to a variety of sensitivity checks. Changing the ordering of the variables left estimates almost unchanged. Similarly, pass-through computed through generalized impulse responses were broadly the same with respect to both size and speed. Estimates computed with different short-term interest rates instead of the monetary aggregate M3 as monetary policy variable also changed little compared to baseline results. Finally, variables were allowed to have long-run relationships and the pass-throughs were estimated via a VECM. Also with this model specification, main conclusions remain unchanged.

The conclusions slightly change, however, when results from the intervention model are considered. Compared to baseline results, the pass-through of exchange rate changes to import prices is weaker in the pre floor regime (around 30 percent in the long run) while substantially stronger in the post floor period (around 70 percent in the long run). Surprisingly, however, the contrary holds for the transmission mechanism of import price shocks to consumer prices, as it appears weaker and only around 12 percent in the post floor period while equally strong, around 25 percent, in the pre floor period. Exchange rate pass-through to consumer prices appears extremely weak for all regimes as in the baseline analysis. This perspective therefore reveals that, while conclusions remain unchanged for the pre floor period, for the post floor period the transmission of exchange rate changes to consumer prices is most likely primarily blocked by a weak transmission mechanism from import to consumer prices rather than by sticky import prices. In light of the fact that the share of imported goods changed little during the period considered, a potential explanation for such results is that domestic producers in competition with imported goods did not adjust their prices in response to the relative price decrease of imported goods following the Swiss Franc appreciation caused by the discontinuation of the floor.

With regards to the floor period, all estimates appear not statistically different from zero. While potentially simply driven by statistical reasons such as sample size or limited data variation, the insignificance of results could also suggest that the introduction of the floor has been effective in stopping the deflationary process characterising the Swiss economy in times proceeding it and creating a substantial pass-through.

In conclusion, findings suggest that the two monetary policy actions considered did affect pass-through dynamics in Switzerland. This provides the need for additional research for mainly two reasons. First, results need to be validated, especially with regards to the effects of the introduction of the floor. Secondly, it is of relevance to assess whether changes in pass-through dynamics caused by the discontinuation of the floor are temporary, reflecting an economy still adjusting to a shock, or whether they represent new equilibrium levels. In addition, further research may be of great value for other small open economies, for which exchange rate pass-through dynamics are likely to be of great importance for price stability.

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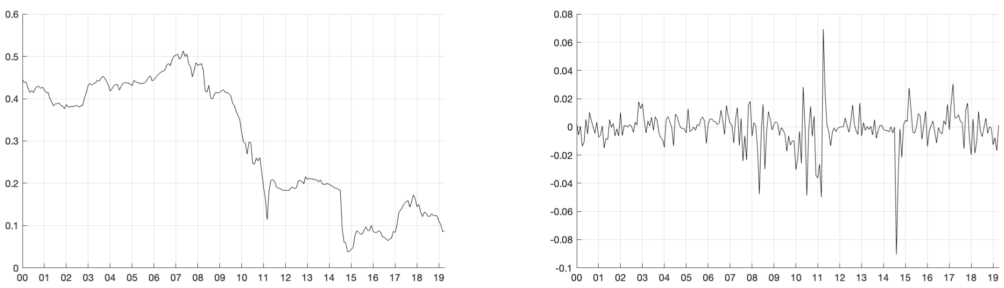
Appendices

A Data

A.1 Descriptive statistics

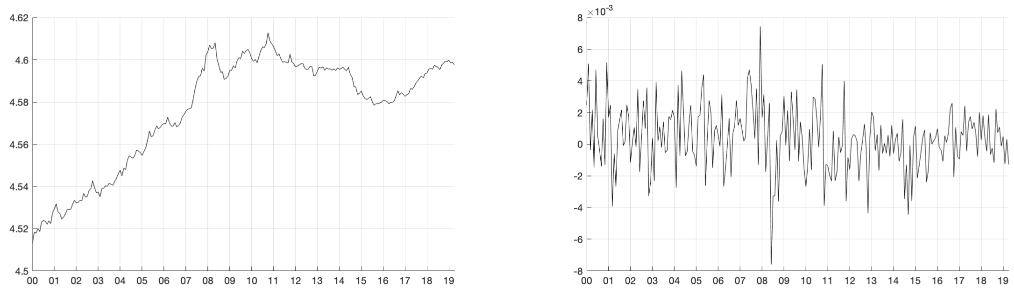
Figure A.1: *EURCHF* Exchange Rate



(a) *EURCHF* Exchange Rate in Log-Levels (b) *EURCHF* Exchange Rate in Log-Differences

Notes: The figure shows monthly EURCHF nominal exchange rate data in log-levels (panel a) and in log-first differences (panel b) for the period 2000:01 to 2019:09. Source: Author's rendering of SNB data (2020).

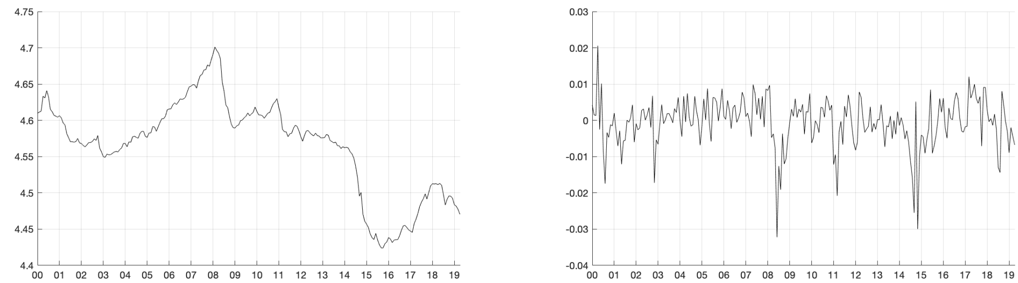
Figure A.2: Swiss Consumer Price Index



(a) Swiss Consumer Price Index in Log-Levels (b) Swiss Consumer Prioce Index in Log-Differences

Notes: The figure shows monthly Swiss Consumer Price Index data in log-levels (panel a) and in log-first differences (panel b) for the period 2000:01 to 2019:09, seasonally adjusted by means of Census-X-12 procedure. Source: Author's rendering of SFSO data (2020).

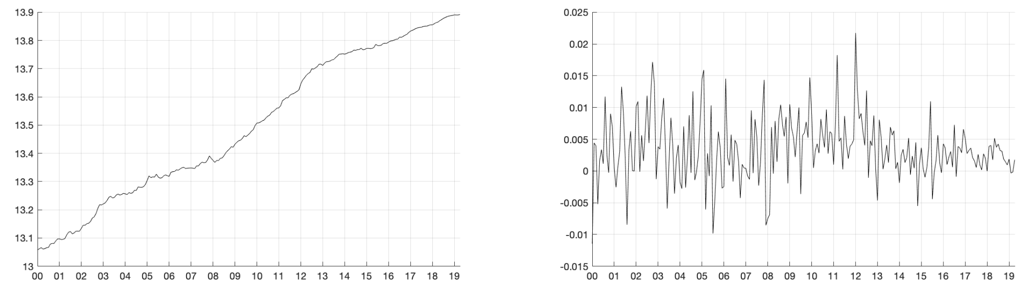
Figure A.3: Swiss Import Price Index



(a) Swiss Import Price Index in Log-Levels (b) Swiss Import Prioce Index in Log-Differences

Notes: The figure shows monthly Swiss Import Price Index data in log-levels (panel a) and in log-first differences (panel b) for the period 2000:01 to 2019:09, seasonally adjusted by means of Census-X-12 procedure. Source: Author's rendering of SFSO data (2020).

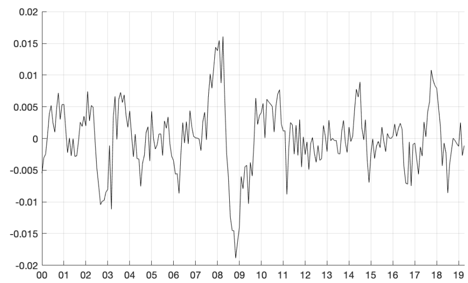
Figure A.4: Monetary Aggregate M3



(a) Monetary Aggregate M3 Log-Levels (b) Monetary Aggregate M3 in Log-Differences

Notes: The figure shows monthly Swiss monetary aggregate M3 data in log-levels (panel a) and in log-first differences (panel b) for the period 2000:01 to 2019:09, seasonally adjusted by means of Census-X-12 procedure. Values in levels are shown in 100 Mio of Swiss Francs. Source: Author's rendering of SNB data (2020).

Figure A.5: Swiss Output Gap



(a) Swiss Output Gap

Notes: The figure shows monthly Swiss output gap values for the period 2000:01 to 2019:09, estimated by means of the Hodrick-Prescott filter method on monthly Swiss GDP data interpolated from observed quarterly Swiss GDP data. Source: Author's rendering of SSEA and SFCA data (2020).

Table A.1: Summary Statistics for Main Variables

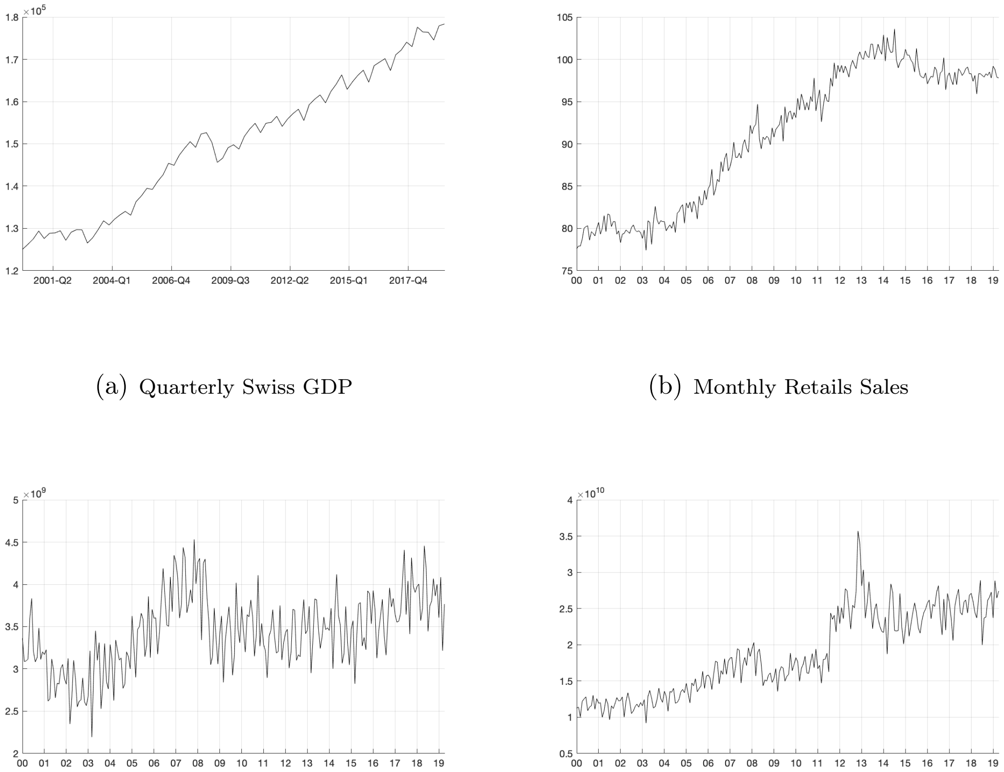
	Variables in levels				**Variables in log-difference**			
Variable	**Mean**	**Std. Dev.**	**Min**	**Max**	**Mean**	**Std. Dev**	**Min**	**Max**
Exchange rate	1.37	0.20	1.04	1.67	-0.0017	0.0137	-0.0906	0.0694
Import Price Index (IPI)	96.39	6.08	83.41	110.09	-0.0005	0.0068	-0.0323	0.0206
Consumer Price Index (CPI)	97.03	2.68	90.98	100.77	0.0004	0.0020	-0.0076	0.0074
Monetary aggregate (M3)	7484.08	1959.43	4686.17	10794.41	0.0035	0.0051	-0.0115	0.0217
Output gap	0.00	0.01	-0.02	0.02	0.0000	0.0053	-0.0189	0.0161
Harmonized CPI	91.62	8.97	74.86	105.41	0.0014	0.0044	-0.0156	0.0134
T	237	237	237	237	236	236	236	236

Notes: The table depicts summary statistics for the main variables used. All variables except for the $EURCH$ nominal exchange rate have been seasonally adjusted by means of the Census-X-12 procedure. M3 shown in 100 Millions when in levels. Sources: Author's rendering of SNB, SFSO, SFCA and SSEA data (2020).

A.2 Estimation of monthly GDP values

Swiss GDP data is available only at quarterly frequency. Therefore, if one wishes to use monthly data, monthly values need to be extrapolated from the quarterly ones. In this study, this is done through the interpolation method proposed by Chow and Lin (1971). Their method consists of using monthly series of variables for which a linear relationship with the true monthly GDP is assumed. The selection of such variables is extremely important for the quality of the monthly estimates. Cuche and Hess (2000), when applying Chow and Lin (1971)'s method to Swiss quarterly GDP data, explain how the series chosen should be i) correlated with the series to interpolate and ii) available at the higher desired frequency. The optimal choice should also be complemented by economic intuition (pp.167-168). When doing a similar exercise, based on these two (three) criteria, Cuche and Hess (2000) propose to base the selection of the related series on the components of the expenditure side of the GDP. However, given the low availability of monthly series for Switzerland, one needs to find good proxies for some of the components. Following both Chow and Lin (1971) and Cuche and Hess (2000), I implement the interpolation using quarterly GDP data and monthly data on retail sales, import of investment goods and exports of all goods, all seasonally adjusted by means of the Census-X-12 procedure. Figure A.6 shows these time series used for the interpolation, while Figure A.7 shows

Figure A.6: Series used for the Interpolation of Quarterly GDP Data



(a) Quarterly Swiss GDP

(b) Monthly Retails Sales

(c) Monthly Imports of Investment Goods

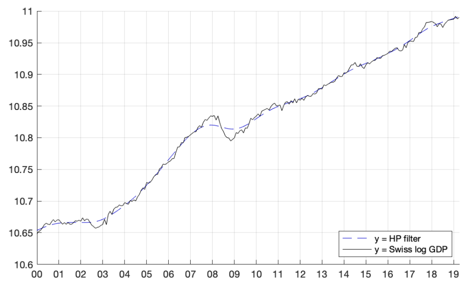
(d) Monthly Exports of all Goods

Notes: The figure depicts quarterly Swiss real GDP values (panel a), monthly retail sales data (panel b), monthly imports of investment goods data (panel c) and monthly exports of all goods data (panel d) for the period 2000:01-2019:09, all seasonally adjusted by means of Census-X-12 procedure. Source: Author's rendering of SSEA and SFCA data (2020).

the resulting monthly GDP estimates together with their trend component computed by means of the Hodrick-Prescott filter method. Cuche and Hess (2000) explain how a good indicator for the quality of the monthly estimates is their volatility compared to the one of the original low frequency series. More precisely, they state how a comparison of the standard deviation of the growth rates is a good indicator. That is, the standard deviation of the growth rate of the monthly series should not be more than four times as large as the one of the quarterly data (p.177). Table A.2 depicts standard deviations.

Despite the standard deviation of the monthly series being slightly lower than that of the quarterly one, I find a ratio of the standard deviations of the growth rates of 1.5, which is well below 4. Furthermore, all coefficients of the estimated Chow and Lin regression are positive. This as well is a good sign for the quality of the estimates, as it implies that the chosen related series are correlated with the quarterly GDP and in the expected direction. Therefore, all in all, monthly estimates appear reasonable.

Figure A.7: Estimated Monthly GDP Series and Hodrick-Prescott Trend Component



(a) Quarterly Swiss GDP

Notes: The figure depicts estimated monthly Swiss real GDP values for the period 2000:01-2019:09, seasonally adjusted by means of Census-X-12 procedure, and its trend component estimated by means of the Hodrick-Prescott filter method. Source: Author's rendering of SSEA and SFCA data (2020).

Table A.2: Standard Deviations of Monthly and Quarterly GDP Series

	Quarterly GDP	**Monthly GDP**
Std. of series in levels	1.6031e+04	5.3227e+03
Std. of growth rate	0.0502	0.0614

Notes: Table depicts standard deviations of the original quarterly GDP series and of the estimated monthly GDP series (both in Mio. of Swiss Francs, real terms) in levels, and the standard deviations of their growth rates. Source: Author's rendering of SSEA and SFCA data (2020).

B Bootstrap confidence intervals for impulse response functions

As explained by Lütkepohl (2005), the validity of asymptotic results can be limited in small samples. For this reason, bootstrapping methods are often preferred to investigate sampling properties of the quantities of interest (p.126). Therefore, the confidence intervals for all the IRF are computed as suggested by Kilian and Lütkepohl (2017) by a standard residual-based recursive design bootstrap method through following steps:

1. Using sample data Y, parameters $\hat{B}$ and $\hat{\Sigma}_u$ and residuals $\hat{U}$ are estimated.
2. Using estimated parameters $\hat{B}$ and residuals $\hat{U}$, a time series of lenght N[1] and denoted by Y^* is simulated. Assuming a stable process, the unconditional mean of the process $(I_{Kp} - \hat{B})^{-1}\nu$ can be used to start the simulation. Residuals are randomly drawn from $\hat{U}$ with replacement. In doing so, one avoids making assumptions about the parametric distribution of the error terms and also ensures that the simulated U^* have the same distribution as $\hat{U}$ (p.342).
3. Using the simulated time series Y^* and residuals U^* computed in the previous step, parameters are re-estimated. These are denoted by $\tilde{B}$ and $\tilde{\Sigma}_u$.
4. Parameters $\tilde{B}$ computed in the previous step are used to compute IRF.
5. Steps 1 to 3 are repeated MC number of times, referred to as the number of Monte Carlo simulations. Here, this has been set to 5000. Results were unchanged when changing this to 10000.
6. Standard percentiles intervals are then computed as $CI_a = [s^*_{\gamma/2}, s^*_{(1-\gamma/2)}]$, where $s^*_{\gamma/2}$ and $s^*_{(1-\gamma/2)}$ denote the $\gamma/2$ and $(1-\gamma/2)$ quantiles of the MC simulated IRF. Here, 95% confidence intervals have been computed, i.e. $\gamma/2 = 0.025$ and $(1-\gamma/2) = 0.975$.

[1] $N = T + B$ where T is the original sample size and B the burn-in period. Here, $B = 100$.

C Lag order selection

Table C.1 and C.2 contain the computed lag order selection criteria for the baseline model as well as the VECM. Numbers in bold correspond to the lag order selected by the criterion in the corresponding column when the maximum number of lags is set to 4 and 8 respectively for the baseline model and to 4 only for the VECM.

Table C.1: Lag Order Selection for Baseline Model

Lags	LR	FPE	AIC	HQ	SC
0	NaN	1.2567e-23	-52.7310	-52.7310	-52.7310
1	287.8801	4.591e-24	-53.738927	**-53.591014**	**-53.371996**
2	70.461597	4.214e-24	-53.825629	-53.529802	-53.091766
3	65.445771	3.953e-24	-53.891077	-53.447337	-52.790283
4	52.424737	**3.921e-24**	**-53.901351**	-53.309699	-52.433626
5	95.979246	3.236e-24	-54.096179	-53.356613	-52.261522
6	67.665922	**3.014e-24**	**-54.171034**	-53.283555	-51.969446
7	39.940431	3.162e-24	-54.128409	-53.093017	-51.559890
8	56.468489	3.096e-24	-54.155818	-52.972512	-51.220367

Notes: Table depicts lag order selection criteria in the following order: sequential Likelihood Ratio (LR) test, Final Prediction Error (FPE), Akaike's Information Criterion (AIC), Hannan-Quinn Information Criterion (HQ) and Schwarz Informatiton Criterion (SC). Critical value for the sequential LR test at the 5% significance level: $\chi^2_{0.95}(25) = 37.652$.

Table C.2: Lag Order Selection for the Vector Error Correction Model

Lags	LR	FPE	AIC	HQ	SC
0	NaN	4.1794e-20	-44.6215	-44.6215	-44.6215
1	4783.0107	**9.7485e-29**	**-64.4992**	**-64.2869**	**-63.9724**
2	**54.7382**	1.0503e-28	-64.4264	-64.0017	-63.3728
3	27.1574	1.2723e-28	-64.2372	-63.6002	-62.6568
4	51.7546	1.3910e-28	-64.1517	-63.3024	-62.0446

Notes: Table depicts lag order selection criteria in the following order: sequential Likelihood Ratio (LR) test, Final Prediction Error (FPE), Akaike's Information Criterion (AIC), Hannan-Quinn Information Criterion (HQ) and Schwarz Informatiton Criterion (SC). Critical value for the sequential LR test at the 5% significance level: $\chi^2_{0.95}(36) = 50.998$

Starting with the sequential LR test, following Lütkepohl (2005), assuming an upper bound M for the lag order, following sequence of null and alternative hypotheses can be tested:

$$\begin{array}{lll} H_0^1 : A_M = 0 & \text{versus} & H_1^1 : A_M \neq 0 \\ H_0^2 : A_{M-1} = 0 & \text{versus} & H_1^2 : A_{M-1} \neq 0 | A_M = 0 \\ & \vdots & \\ H_0^i : A_{M-i+1} = 0 & \text{versus} & H_1^i : A_{M-i+1} \neq 0 | A_M = ... = A_{M-i+2} = 0 \\ & \vdots & \\ H_0^M : A_1 = 0 & \text{versus} & H_1^M : A_1 \neq 0 | A_M = ... = A_2 = 0 \end{array}$$

That is, each null-hypothesis is tested conditionally on the previous ones being true. The VAR order is then chosen according to the first null hypothesis that can be rejected: if H_0^i is rejected, the optimal lag order corresponds to $p = M - i + 1$ (pp.143-144). The LR test value is computed as:

$$\lambda_{LR}(i) = T[\log|\hat{\Sigma}_u(M-i)| - \log|\hat{\Sigma}_u(M-i+1)|]$$

where $\hat{\Sigma}_u(m)$ denotes the ML estimator of Σ_u when a VAR(m) model is fitted to a time series of length T. The statistic has a $\chi^2(K^2)$ asymptotic distribution if H_i^0 and all previous null hypotheses hold.

The final prediction error (FPE) criteria is based on minimizing the forecast mean squared errors and is computed in the following way:

$$FPE(m) = \left[\frac{T+km+1}{T-km-1}\right] det\hat{\Sigma}_u(m)$$

where $m = 0, ..., M$ is the order of the estimated model, k the number of endogenous variables and T the sample size. Then, the order minimizing the FPE values should be chosen (Lütkepohl, 2005, pp.146-147).

Similarly, for the remaining criteria, the lag order to be chosen is the one that minimizes them. They are computed as follows:

$$\begin{aligned} AIC(m) &= ln|\hat{\Sigma}_u(m)| + \frac{2}{T}(mk^2) \\ HQ(m) &= ln|\hat{\Sigma}_u(m)| + \frac{2\text{loglog}T}{T}(mk^2) \\ SC(m) &= ln|\hat{\Sigma}_u(m)| + \frac{\log(T)}{T}(mk^2) \end{aligned}$$

See, again, Lütkepohl (2005) for more details (pp.147-151). As he explains, it is known that, in small samples, AIC and FPE may have better properties, i.e. choose the correct order more often, than HQ and SC.

To check if residuals are well behaved, Portmanteau tests have been implemented. A Portmanteau test tests the null hypothesis of no autocorrelation in the residuals up to lag h against the alternative of some autocorrelation being present. Formally:

$$H_0 : R_h = (R_1, ..., R_h) = 0 \text{ versus } H_1 : R_h \neq 0$$

where R_i represents the autocorrelation matrix of residuals u_t for lag i. Then, the test value is computed as:

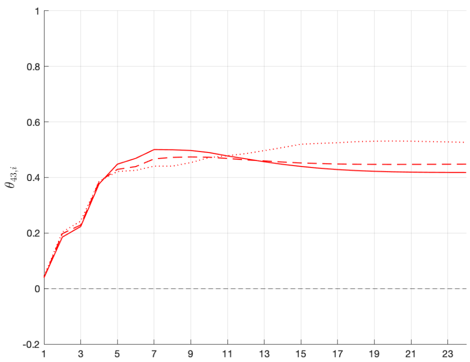
$$Q_h = T^2 \sum_{i=1}^{h} (T-i)^{-1} tr(\hat{C}_i' \hat{C}_0^{-1} \hat{C}_i \hat{C}_0^{-1})$$

where $\hat{C}_i$ represents the estimated autocovariance matrix of residuals u_t at lag i and T the sample size. The test value has an approximate asymptotic χ^2 distribution with $K^2(h-p)$ degrees of freedom, where K is the number of endogenous variables, h the maximal lag length of the test and p the VAR order. For more details on the computation of the test see Lütkepohl (2005).

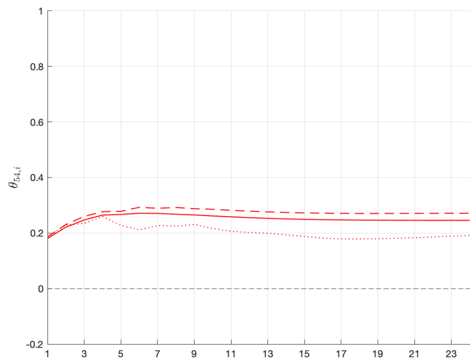
For the estimated VARX(3, 3) model, the computed Q_{18} test value equals 397.420, while the corresponding critical value at the 1% significance level is given by $\chi^2_{0.99}(K^2(h-3)) =$ 441.634. Since Q_{18} is smaller than its critical value, the null hypothesis cannot be rejected. This is taken as suggestive evidence for well behaved residuals.

Given the uncertainty surrounding the choice of the lag order, it is of interest to investigate the sensitivity of results with respect to such choice. Figure C.1 shows main estimates of the baseline specification for the chosen order of 3 lags and other two orders (6 and 4 lags respectively). As can be seen from the picture, results are broadly the same with all orders.

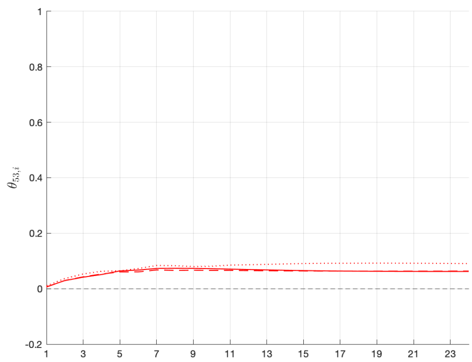
Figure C.1: Accumulated Impulse Response Functions under Different Lag Orders



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices

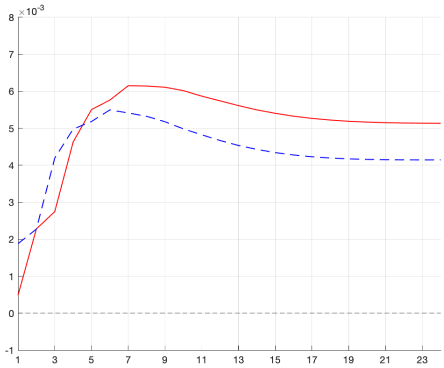


(c) ERPT to Consumer Prices

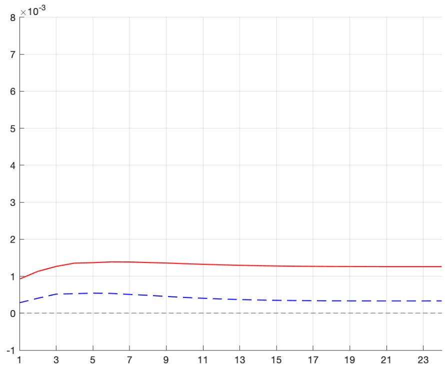
Notes: The figure shows accumulated orthogonalised impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a one-unit shock in the impulse variable based on the baseline VARX(3,3) model. Dashed lines and dotted lines represent accumulated orthogonalized impulse response functions of a one-unit shock in the impulse variable based on the same model, but with 4 and 6 lags respectively for both endogenous and exogenous variables.

D Generalized impulse response functions

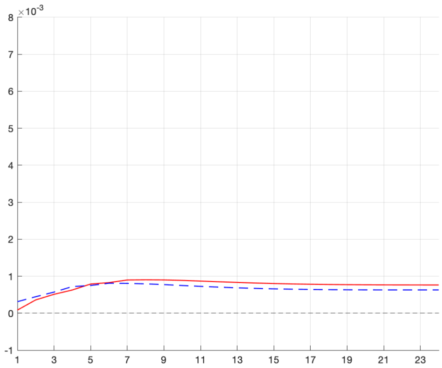
Figure D.1: Accumulated Generalized Impulse Response Functions



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices



(c) ERPT to Consumer Prices

Notes: The figure shows accumulated impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a standard deviation shock in the impulse variable based on the baseline VARX(3,3) model. Blue dashed lines represent accumulated generalized impulse response functions of a standard deviation shock in the impulse variable based on the same model.

E Bai and Perron testing procedure

Here, the Bai and Perron (1998) testing procedure is explained in more detail. Formally, consider following multivariate least square regression model:

$$y_t = x_t'\beta + z_t'\delta_j + u_t \tag{E.1}$$

for $t = T_{j-1} + 1, ..., T_j$ and $j = 1, ..., m$ with $m + 1$ being the number of regimes (and m the number of breaks), y_t is the observed independent variable, x_t with dimension $(p \times 1)$ and z_t with dimension $(q \times 1)$ are vectors of covariates, and, β and δ_j $(j = 1, ..., m+1)$ are the corresponding vectors of coefficients; u_t is the error term. The indices $(T_1, ...T_j)$, or the break points, are treated as unknown. The idea is to estimate the unknown regression coefficients together with the break points when T observations on (Y_t, x_t, Z_t) are available by means of least-square. Written as in equation E.1, the model is one of a partial structural break in the sense that β is not subject to changes and is estimated using the entire sample. When $p = 0$ (i.e. when all regressors are included in z_t), a pure structural break model is obtained, i.e. one where all the coefficients are subject to a change. With regards to the error terms u_t, the only restriction imposed is that if lagged variables are allowed in $\{x_t, z_t\}$, no serial correlation is permitted in $\{u_t\}$. On the contrary, if no lagged variables are included, the error terms are allowed to be serially correlated.

The implementation of the Bai and Perron testing procedure is quite demanding, since it requires a high number of estimations. The authors, however, developed an algorithm based on dynamic programming which limits the amount of least-squares operation to a certain order. This study used the Matlab code put at disposal by the same authors[1] to implement such algorithm through which the testing procedure can be carried out. When using it, for all series, a linear regression model in the form of an AR process has been specified:

$$y_t = \mu + \sum_{i=1}^{p} \rho y_{t-i} + u_t$$

For the Δe_t series, one lag has been included according to the BIC criteria. For both Δcpi_t and Δipi_t, the BIC criteria selects two lags. Since all linear regression include lagged variables as regressors, no correlation in the residuals has been allowed for. The algorithm requires to specify regressor matrices x_t and z_t, so to determine whether to test

[1] The Matlab code can be retrieved from Perron's personal website: http://people.bu.edu/perron/code.html

for a pure or a partial structural break. When testing for partial breaks, constant μ only has been included in z_t, while the lagged variables in x_t. When pure structural breaks are tested for, all parameters have been included in z_t (and, accordingly, x_t has been set to 0). The algorithm further requires the specification of a so called trimming parameter, which represents the minimal length of a sample segment (observations between two breaks), namely $\varepsilon = h/T$, where h is the minimum number of observations, and whether one wants to allow for heteroskedasticity in the error terms or not. Following the common practice, a trimming parameter of 0.15 percent has been imposed for all tests. Error terms have been allowed to have varying variance for all tested series. Finally, the maximal number of possible breaks considered in the double maximum test was set to 5.

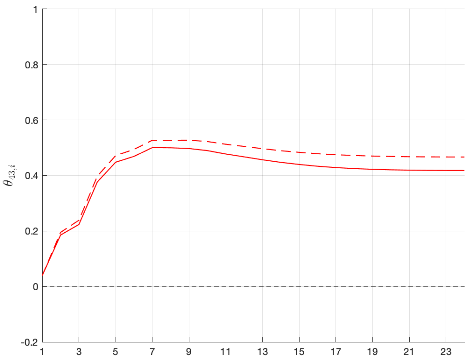
In detail, results are the following. For the exchange rate series, as mentioned in Section 6, no breaks have been identified. More precisely, when partial breaks are tested for, the sequential procedure selects 0 breaks and the null hypothesis of the double maximum test cannot be rejected at neither the 10% nor the 5% level, as the computed test value of 4.42 is smaller than both critical values 7.46 (10% significance level) and 8.88 (5% significance level). BIC and LWZ also select 0 breaks. With regards to pure structural breaks, same conclusions hold. The double maximum test value is 5.859, which is smaller than both 10% and 5% significance level critical values of 10.16 and 11.70 respectively.

For the CPI series, when partial breaks are tested for, the sequential testing procedure estimates one break in August 2008 at the 2.5% significance level. While not by the LWZ, results are confirmed by the BIC, which also identifies one break. In contrast, no pure structural breaks are estimated. The double maximum test value amounts to 11.08, which is smaller than both 12.40 and 14.23, the critical values at 10 and 5% significance level, indicating that the null hypothesis of no structural breaks cannot be rejected.

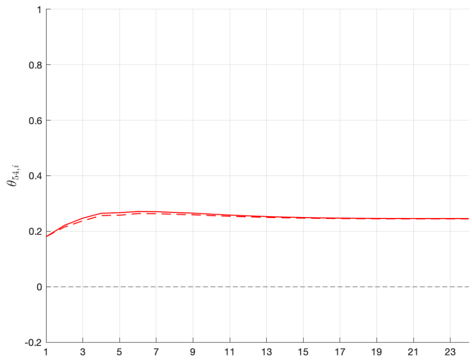
Finally, for the IPI series, no partial break is estimated. Double maximum test value is 4.76, while critical values at 10 and 5% significance levels amount to 7.46 and 8.88 respectively. Results are confirmed by both BIC and LWZ criteria. However, two pure structural breaks are identified. The sequential procedure is not able to reject the null hypothesis of zero breaks against one. However, with a test value of 14.27 and 10 and 5% significance level critical values of 12.4 and 14.23, the null hypothesis of no breaks implied by the double maximum test can be rejected. Therefore, going back to the sequential procedure reveals that, indeed, two breaks are estimated. The first is estimated in August 2008 and the second in April 2015.

F Controlling for the financial crisis

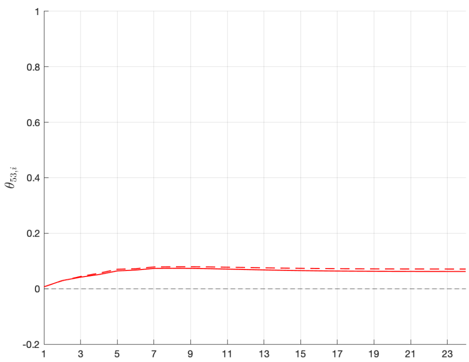
Figure F.1: Controlling for the Financial Crisis: Accumulated Impulse Response Functions



(a) ERPT to Import Prices



(b) Pass-Through of Import to Consumer Prices



(c) ERPT to Consumer Prices

Notes: The figure shows accumulated orthogonalised impulse response functions. Red solid lines represent accumulated orthogonalised impulse responses to a one-unit shock in the impulse variable based on the baseline VARX(3,3) model. Dashed lines represent accumulated orthogonalized impulse response functions of a one-unit shock in the impulse variable based on the same model, but including a dummy variable to control for the Financial crisis of 2008.

G Intervention model estimation

G.1 Maximum Likelihood estimation under different hypotheses

As explained in the main text, ML estimation of the intervention model implied by equation 7.1 becomes fairly trivial when certain restrictions are imposed on the parameters of the model. Below the ML estimators for B and Σ_{ut} under each of the hypotheses tested, together with the respective test values (Lütkepohl, 2005, pp.595-601):

H_1^0 : stationary model with constant parameters. Formally:

$$H_0^1\text{: } B_i = B_1 \text{ and } \Sigma_{ui} = \Sigma_{u1} \text{ for } i = 2, 3$$

ML estimators are computed as:

$$\tilde{B}_1^{(1)} = \left(\sum_t y_t Z_{t-1}'\right)\left(\sum_t Z_{t-1}Z_{t-1}'\right)^{-1}$$
$$\tilde{\Sigma}_{u1}^{(1)} = \sum_t (y_t - \tilde{B}_1^{(1)} Z_{t-1})(y_t - \tilde{B}_1^{(1)} Z_{t-1})'/T$$

where now $Z_{t-1} = (1, Y_{t-1}')'$ and T is total sample size. The test value to be used when testing this hypothesis against an alternative corresponds to

$$\lambda_0^1 = -\frac{1}{2} T \log|\Sigma_{u1}^{(1)}|$$

H_0^2 : time varying intercept, constant coefficients and variance. Formally:

$$H_0^2\text{: } \nu_t = \textstyle\sum_{i=1}^3 n_{it}\nu_i;\ A_i = A_1 \text{ and } \Sigma_{ui} = \Sigma_{u1} \text{ for } i = 2, 3$$

ML estimators are obtained by first defining following matrices:

$$W_{t-1} = \begin{bmatrix} n_{1t} \\ n_{2t} \\ n_{3t} \\ Y_{t-1} \end{bmatrix}; A = \begin{bmatrix} \nu_1, \nu_2, \nu_3, A_1 \end{bmatrix}$$

Then, ML estimators are computed as

$$\tilde{C} = \left(\sum_t y_t W'_{t-1}\right)\left(\sum_t W_{t-1}W'_{t-1}\right)^{-1}$$
$$\tilde{\Sigma}_{u1}^{(2)} = \sum_t (y_t - \tilde{C}W_{t-1})(y_t - \tilde{C}W_{t-1})'/T$$

and the corresponding test value as

$$\lambda_0^2 = -\frac{1}{2}T\ \log|\tilde{\Sigma}_{u1}^{(2)}|$$

H_0^3: time varying intercept and coefficients, constant variance. Formally:

$$H_0^3\text{: } B_t = [\nu_t, A_1] = \textstyle\sum_{i=1}^3 n_{it}B_i \text{ and } \Sigma_{ui} = \Sigma_{u1} \text{ for } i = 2, 3$$

ML estimators are computed as

$$\tilde{B}_i^{(3)} = \left(\sum_t n_{it} y_t Z'_{t-1}\right)\left(\sum_t n_{it} Z_{t-1}Z'_{t-1}\right)^{-1}$$
$$\tilde{\Sigma}_{u1}^{(3)} = \sum_{i=1}^{3}\sum_t n_{it}(y_t - \tilde{B}_i^{(3)}Z_{t-1})(y_t - \tilde{B}_i^{(3)}Z_{t-1})'/T$$

and the corresponding test value as

$$\lambda_0^3 = -\frac{1}{2}T\ \log|\tilde{\Sigma}_{u1}^{(3)}|$$

H^a : all time varying parameters. Formally:

$$H^a\text{: } B_t = [\nu_t, A_1] = \textstyle\sum_{i=1}^3 n_{it}B_i \text{ and } \Sigma_{ut} = \sum_{i=1}^3 n_{it}\Sigma_{ui}$$

ML estimators are computed as

$$\tilde{B}_i^{(a)} = \tilde{B}_i^{(3)}$$
$$\tilde{\Sigma}_{ui}^{(a)} = \sum_t n_{it}(y_t - \tilde{B}_i^{(a)}Z_{t-1})(y_t - \tilde{B}_i^{(a)}Z_{t-1})'/(T\bar{n}_i)$$

where $\bar{n}_i = (1/T)\sum_t n_{it}$ and the corresponding test value

$$\lambda_a = -\frac{1}{2}T(\bar{n}_1\ \log|\tilde{\Sigma}_{u1}^{(a)}| + \bar{n}_2\ \log|\tilde{\Sigma}_{u2}^{(a)}| + \bar{n}_3\ \log|\tilde{\Sigma}_{u3}^{(a)}|)$$

G.2 Chow tests computation

As explained in the main text, when testing all the null hypotheses against the single alternative hypothesis, all tests take the form of LR tests.

LR tests are computed through the following general formula:

$$\lambda_{LR} = 2(\lambda_a - \lambda_0)$$

where λ_a and λ_0 stand for the test value under the alternative and under the null hypothesis respectively. For more details on the computational aspects of the LR tests see Lütkepohl (2005).

G.3 Bootstrap tests computation

As explained by **?**, the aim of bootstrap tests is to estimate the distribution of a test statistic under the DGP that generated it, provided that the DGP satisfies the null hypothesis (p.5). If the distribution of the test values under the null hypothesis also depends on some nuisance parameters[1] then one should compute parametric bootstrap tests (Davison & Hinkley, 1997, p.148). The procedure is similar to the one implemented for the bootstrap confidence intervals of the IRF. MC independent replications of the sample $y_1, ..., y_T$ are drawn from the model under the null hypothesis. In other words, MC time series of length T are simulated using parameters estimated under the null hypothesis. For each of the MC-th sampled time series, the test statistic is computed. After all MC test statistics have been computed, the parametric p-value is calculated as:

$$p_{bootstrap} = \frac{1 + \#\{t_{mc} \geq t\}}{MC + 1}$$

where t_{mc} represents the test statistic computed with the MC-th simulated time series and t the theoretical asymptotic test value computed as described in Section G.2. In other words, the bootstrap p-value represents the share of the simulated test statistics that are larger than the asymptotic test value. For more details on the bootstrapping method see Davison and Hinkley (1997).

[1] A nuisance parameter refers to any parameter which is not of immediate interest but which must be accounted for in the analysis of those parameters which are of interest.