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Identifying US Monetary Policy Shocks through Sign Restrictions in Dollarized Countries

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Abstract

Since dollarized countries import US monetary policy, identifying US monetary shocks through sign restrictions on US variables only, does not use all available information. In this paper we therefore include dollarized countries, which enable us to restrict more variables and leave the responses of US output *and prices* unrestricted (to allow for the working capital view of monetary shocks). We find only little evidence for the latter in the US, as prices fall immediately after most contractionary shocks that we identify. Furthermore, monetary shocks do not seem to have a clear effect on real GDP.

JEL-classification: E52; E31; C32

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1 Introduction

Following the seminal work by Faust (1998), Canova and De Nicoló (2002) and Uhlig (2005), the use of sign restrictions has become a popular way to identify structural

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shocks in the vector autoregressive (henceforth: VAR) literature. Although this approach has much intuitive appeal, it has not yet managed to bring consensus in the debate on what the effects of monetary policy shocks are. Canova and De Nicoló (2002) for example find that contractionary monetary shocks lead to a decrease in output, while Uhlig's (2005) agnostic identification procedure ("agnostic" since he leaves the response of output unrestricted) fails to detect a clear effect. Actually, Uhlig (2005) even finds that GDP tends to go *up* in response to the majority of the contractionary shocks that satisfy the imposed restrictions (see his Figure 3). As this is inconsistent with all theoretical models we know of, one could dub this an "output puzzle".

The reason for this surprising result may lie in the fact that a number of different shocks can satisfy a certain set of restrictions. For example, with sign restrictions on the responses of US prices and monetary variables (as in Uhlig (2005)), it is possible that one also includes a positive technology shock in, what is supposed to be, a contractionary monetary shock. After all, if one wishes to stay agnostic with respect to the sign of the output response, a contractionary monetary and a positive technology shock can only be distinguished by the response of monetary variables in a standard monetary VAR-specification.[1] But if the monetary authority is not able to recognize technology shocks in real time (which seems plausible given that researchers cannot even agree on what they look like *ex post*; compare Galí (1999) with Christiano, Eichenbaum and Vigfusson (2003)), the identifying restriction on these monetary instruments is not that useful anymore and the procedure might confuse the two shocks, possibly contributing to the output puzzle.

In addition, Paustian (2007) shows that the agnostic identification procedure is only able to uncover the correct sign of the unconstrained response if one restricts a large enough number of variables.

In this paper, we try to address these two concerns by including price data from four dollarized countries, all located in Latin America. We argue that this approach has two advantages. First, it simply exploits the fact that these countries contain useful information for the identification of monetary US shocks: dollarized countries also use the US dollar as legal tender (just like Ohio or any other US state does) and hence import the same Federal Reserve policy as normal US states do. Consequently, US monetary policy should not only affect the relevant US variables, but also those in the client countries. Once one recognizes this, one has more variables to place sign restrictions upon, which should help to identify the monetary shocks more precisely.

[1] Both shocks are supposed to decrease the price level, but after a positive technology shock monetary policy should respond in an expansive way, which distinguishes it from a contractionary monetary shock (see for example Table 2 in Paustian (2007)).

Second, including data from other countries makes the procedure less vulnerable to the possible confusion of shocks. This stems from the fact that the US economy is not perfectly integrated with those of the dollarized countries (see the discussion in Willems (2011)). Consequently, the transmission of non-monetary US shocks to the client countries is anything but perfect.[2] In particular, a non-monetary US shock is unlikely to affect prices in all dollarized countries contemporaneously, as these shocks generally need some time to be transmitted. Monetary shocks, on the other hand, are transmitted rapidly through financial markets. This gives the sign restriction procedure an additional criterion to distinguish a monetary US shock from a non-monetary one: the former does affect the price level in the dollarized countries at short horizons, while the latter is unlikely to do so.[3]

Next to this, it is also possible that the traditional sign restriction procedure misidentifies monetary shocks because the latter do not lead to an *immediate* drop in prices. This result actually shows up quite often in non-sign restricted VAR-exercises, where it is known as the "price puzzle" (*cf.* Sims (1992)). In fact, the sample period considered in this paper also contains a price puzzle if one uses the popular Choleski scheme to identify the monetary shock (see the Appendix).

The initial price increase after a monetary contraction is seen as "puzzling" as it is inconsistent with most theoretical models used by economists, such as the standard New Keynesian one. It is however perfectly consistent with models in which cost-side considerations play a role: if firms need to borrow working capital in order to be able to pay for their production factors, the interest rate becomes a determinant of real marginal costs and prices will show a short-lived increase after a contractionary monetary shock.[4] However, studies employing sign restrictions typically do not allow for the working capital view of monetary shocks, as they tend to define contractionary monetary shocks as shocks that (among other things) decrease the

[2] In fact, Canova (2005) finds the even stronger result that non-monetary US shocks do not produce significant output and price responses in Latin American countries at all.

[3] Similarly, a non-monetary shock emanating from any of the dollarized countries is unlikely to affect prices in all other countries that use the US dollar (especially at short horizons). This makes them distinguishable from a US monetary shock as well. Our procedure would have difficulties with an aggregate world-wide technology shock, as that would affect prices in all included countries instantaneously in the same direction as a US monetary shock would. But given the evidence in Gabaix (2011), who shows that many "aggregate US shocks" are actually idiosyncratic shocks to large US firms, one can question how common "aggregate world-wide technology shocks" are.

[4] See Van Wijnbergen (1983), who obtained the price puzzle in such a model *avant la lettre*, and Barth and Ramey (2001), who call for a more serious consideration of the cost channel; Christiano, Eichenbaum and Evans (2005) come up to Barth and Ramey's call, as they construct a DSGE-model which is able to replicate the price puzzle through the working capital channel. Ravenna and Walsh (2006) discuss the consequences of the cost channel for optimal monetary policy.

price level immediately.

Given the importance of understanding how prices respond to policy shocks (after all, the price response determines whether a certain action will manage to stabilize the economy or not), this paper tries to remain more agnostic on this issue: in our benchmark specification, we identify US monetary shocks by placing sign restrictions on the federal funds rate and on prices *in the dollarized countries only.* So while Uhlig's (2005) procedure only leaves the response of US output unrestricted (and avoids the US price puzzle by construction), we leave both US output *and prices* unrestricted. In addition, we only impose the negativity restrictions on prices in the dollarized countries after several months, to allow those price levels to show a short-lived price increase following a monetary contraction (as would be the case if the cost channel plays a non-negligible role).

The results of this exercise however strongly suggest that US prices fall immediately after a monetary contraction. Despite our agnosticism on this issue, we hardly find any shocks that satisfy the working capital view on monetary policy in the US. In the dollarized countries we find more shocks that are consistent with this view, suggesting that the working capital channel is somewhat more important for emerging economies than it is for the US.

Additionally, we fail to find a clear effect of US monetary shocks on US output, so we cannot reject monetary neutrality. This finding is robust to restricting the response of US prices, while restricting the response of US output shows that it is quite difficult to find shocks that satisfy the standard New Keynesian view of a monetary policy shock.

2 Approach

We start with the estimation of a reduced form VAR with p lags for a vector of variables Z_t, which is of size $(m \times 1)$:

$$Z_t = \sum_{i=1}^{p} B_i Z_{t-i} + e_t \tag{1}$$

In this equation, $Z_t = \left(Z_t^{US}, Z_t^{D1}, Z_t^{D2}, ..., Z_t^{Dn}\right)'$, where Z_t^{US} is a $(m_1 \times 1)$-vector containing US-variables and Z_t^{Dj} is a $(m_2 \times 1)$-vector with data from the j^{th} dollarized country. The B_i's are the coefficient matrices and e_t is the reduced form error with variance-covariance matrix Σ.[5]

[5]A constant term that is also included in the analysis is omitted here for notational simplicity.

When estimating this system, we impose some additional structure on the B-matrices. First, we follow (among others) Cushman and Zha (1997) and assume that the US economy is exogenous with respect to the small economies of the dollarized countries. That is: whatever happens in the dollarized countries is assumed to have no impact on the US variables, as the former are assumed to be too small to affect the latter. Second, the variables in all dollarized countries are assumed to be independent of each other, conditional on the US variables. So whatever happens in dollarized country h is assumed to have no impact on the variables in dollarized country $j \neq h$. This assumption keeps the VAR parsimonious (it makes the south-east block of all B-matrices diagonal), but has the drawback that we probably miss some of the interdependencies between the dollarized countries. In our view, the latter are however unlikely to play an important role for the question we are interested in, so we disregard them in our empirical approach.

Econometrically, the stated assumptions boil down to estimating the following system:

$$\begin{bmatrix} Z_t^{US} \\ Z_t^{D1} \\ Z_t^{D2} \\ \vdots \\ Z_t^{Dn} \end{bmatrix} = \sum_{i=1}^{p} \begin{bmatrix} B_{11,i} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} \\ B_{21,i} & B_{22,i} & \mathbf{0} & \cdots & \mathbf{0} \\ B_{31,i} & \mathbf{0} & B_{33,i} & \cdots & \mathbf{0} \\ \vdots & \vdots & \ddots & \ddots & \vdots \\ B_{(n+1)1,i} & \mathbf{0} & \mathbf{0} & \cdots & B_{(n+1)(n+1),i} \end{bmatrix} \begin{bmatrix} Z_{t-i}^{US} \\ Z_{t-i}^{D1} \\ Z_{t-i}^{D2} \\ \vdots \\ Z_{t-i}^{Dn} \end{bmatrix} + e_t \qquad (2)$$

It is apparent that regressors differ per equation, thus making OLS estimates inefficient. We therefore estimate the whole system using the seemingly unrelated regressions procedure (*cf.* Zellner (1962)).

We identify shocks by imposing sign restrictions on the impulse responses. Since a structural VAR with sign restrictions is not exactly identified (see Fry and Pagan (2007)), we look for the set of structural VARs that satisfy the restrictions while sharing the same, estimated reduced form representation. To do so, we first invert the VAR to get the VMA(∞) representation:

$$Z_t = C(L)u_t$$

Here, $u_t = P^{-1}e_t$ are orthonormal shocks, $C(L) = (I - B(L))^{-1}P$, and P is the lower triangular Choleski decomposition of Σ (such that $P'P = \Sigma$). The coefficients of the matrix polynomial $C(L) = C_0 + C_1L + \ldots$ correspond to the orthogonalized impulse response functions, which, in general, do not satisfy the desired sign restrictions. The idea of the sign restrictions procedure then is to generate a large number of orthogonal matrices $Q^{(i)}$ and check whether each alternative structural representation

$$Z_t = C(L)Q^{(i)\prime}Q^{(i)}u_t = C(L)Q^{(i)\prime}\eta_t^{(i)} = D^{(i)}(L)\eta_t^{(i)}$$

corresponds to impulse responses that are consistent with the sign restrictions.[6]

To generate the $Q^{(i)}$-matrices we use the algorithm of Rubio-Ramirez, Waggoner and Zha (2010). This algorithm exploits the fact that for a given square matrix M, its QR decomposition $M = QR$ is formed by an orthogonal matrix Q and a triangular matrix R. Hence, we can get one orthogonal matrix $Q^{(i)}$ after drawing a random normal square matrix $M^{(i)}$.

Note that we only place restrictions on the responses after an impulse to the equation for the monetary policy instrument in Z_t^{US}. If the restrictions are satisfied for all K periods, we keep the structural representation $D^{(i)}(L)$ and label the corresponding shock $\eta_{r,t}^{(i)}$ (where r is the position of the monetary policy instrument in Z_t^{US}) as a "monetary policy shock". If not, the candidate shock is discarded.

We report the resulting set of impulse responses following the pure-sign-restriction approach set out in Uhlig (2005), so the interested reader is referred to that source for all remaining details.

3 Data and VAR specification

We have been able to obtain data on prices for four dollarized countries: Ecuador, El Salvador, Panama and Puerto Rico.[7] The major obstacle to the practical implementation of our approach, is the limited length of some of the time series used. Although Panama and Puerto Rico have been dollarized for over a hundred years already (with data available since 1974 and 1989, respectively), we are constrained by the fact that Ecuador became officially dollarized in 2000, while El Salvador only introduced the US dollar in 2001.[8] Therefore, we keep the VAR parsimonious and

[6] Note that since the $Q^{(i)}$-matrices are such that $Q^{(i)\prime}Q^{(i)} = Q^{(i)}Q^{(i)\prime} = I$, the rotated shocks $\eta_t^{(i)}$ are orthogonal with unit variance, as $E(\eta_t^{(i)}\eta_t^{(i)\prime}) = Q^{(i)}P^{-1}E\left(e_t e_t'\right)P^{-1\prime}Q^{(i)\prime} = Q^{(i)}P^{-1}PP'P^{-1\prime}Q^{(i)\prime} = I$.

[7] Although we will sometimes refer to it as "country", one should note that Puerto Rico is actually not a country, but a US associated free state.

[8] However, El Salvador's former currency (the colón) was already pegged to the US dollar since 1993. Similarly, Ecuador was already effectively dollarized in the mid-1990s, by which time its residents had switched to using US dollars in their daily transactions and had the majority of their deposits/loans denominated in dollars (Beckerman, 2001). So in practice both Ecuador and El Salvador had already been importing US monetary policy for some years, when they dollarized officially in the early 2000s.

only include the essential variables in our benchmark specification. In particular, we use data (taken from the St Louis Fed website) on US real GDP, the GDP deflator, the federal funds rate, and prices in the four dollarized economies considered. However, as will be shown in Section 5 of this paper, results are robust to including other variables in the VAR (such as reserves or monetary aggregates).

Following Uhlig (2005), we estimate the VAR on monthly observations. As the series for GDP and the GDP deflator are only available at the quarterly frequency, we interpolate these variables using the method described in Chow and Lin (1971). For GDP, we do this using the industrial production index, while the deflator is interpolated with help of the CPI and PPI. For the dollarized countries, the GDP deflator is not available at all, as a result of which we work with the monthly series of the local CPI's instead, taken from the IMF's IFS database.[9]

To reduce the risk of misspecification, we estimate the VAR on post-1989:1 data. This is motivated by the findings of Bagliano and Favero (1998), who report evidence of a structural break in the US monetary policy rule in 1988 (just after Greenspan succeeded Volcker as chairman of the Board of Governors of the Federal Reserve). We have also estimated the VAR on samples with different starting dates (both earlier and later), but found nearly identical results (see Section 5.2).

Since data for the dollarized economies are relatively scarce, we exploit all available time series up to their full length. This implies that we estimate the VAR-coefficients for the US, Panama and Puerto Rico on a sample running from 1989:1 to 2010:7,[10] while the coefficients for Ecuador and El Salvador are estimated on samples running from 2000:1 to 2010:7 and 2001:1 to 2010:7, respectively (as these countries have only been officially dollarized since the early 2000s).[11] In our estimation we account for the unequal number of observations among countries by using the "usual" estimator for the variance-covariace matrix (see Schmidt (1977)). All variables are logged before they enter the VAR (except for the federal funds rate, which enters in levels). Finally, we chose the number of lags for each country block based on

[9] Although there also exist quarterly data on real GDP for Ecuador, El Salvador and Panama (not for Puerto Rico), we do not include them here for two reasons. First, one can question the usefulness of these data if one wishes to stay agnostic with respect to the response of output (especially if one wants to keep the VAR parsimonious). Second, there are no monthly indicators of real activity available for these countries, as a result of which it is difficult to obtain imputed series for GDP at the monthly frequency.

[10] Recall that the starting date follows Bagliano and Favero (1998).

[11] See Canova (2005) for a similar approach. In the Appendix we show that restricting the analysis to the US, Panama and Puerto Rico (so that all coefficients can be estimated on data covering the same sample period from 1989:1 to 2010:7) does not significantly affect the results. In addition, our findings are also robust to restricting the analysis to the much shorter post-2001 sample, during which all countries considered in this paper were officially dollarized.

Akaike's information criterion, which results in selecting three lags for the US, five for Ecuador, one for El Salvador, three for Panama and six lags for Puerto Rico.

4 Results

This section describes the findings obtained by applying sign restrictions to a VAR which also includes variables from dollarized countries. As argued in the Introduction, this allows us to restrict a larger number of variables, which should help to identify the US monetary policy shock more precisely (Paustian, 2007).

In addition, the economies of the US on the one hand, and those of the dollarized countries on the other, are only imperfectly integrated with each other (*cf.* Willems (2011)). This gives us an additional criterion to distinguish monetary US shocks from non-monetary ones, as the former are transmitted rapidly to the client countries via financial markets, while the transmission of non-monetary US shocks to these countries is far from perfect (Canova, 2005). In particular, non-monetary shocks typically need more time to be transmitted, as this mainly occurs through the time consuming trade channel.[12] Consequently, they are unlikely to induce responses of variables in dollarized countries at short horizons, which distinguishes them from monetary US shocks.

As emphasized before, we also try to acknowledge the idea that the price "puzzle" might actually be a fact, as it could just be caused by the working capital channel. Consistent with this mechanism, we therefore allow for a short-lived price increase after a contractionary shock by imposing the restrictions on prices only as of the fourth month after the shock has hit.[13]

Section	4.1	4.2	4.3
US GDP			≤ 0
US GDP deflator		$\leq 0^*$	≤ 0
Federal funds rate	≥ 0	≥ 0	≥ 0
Dollarized CPIs ($4\times$)	$\leq 0^*$	$\leq 0^*$	≤ 0

Table 1: Summary of imposed sign restrictions in different sections. An asterisk (*) indicates that this restriction is only imposed as of the fourth month after the shock has hit.

[12]This observation has a long standing tradition in international economics, going back to at least Dornbusch (1976).

[13]Considering different starting periods for the sign restrictions on prices produces very similar results (see the Appendix).

Table 1 summarizes the different sets of sign restrictions we will impose to try to identify the US monetary policy shocks with. In our benchmark specification, we impose the restrictions until one year after the shock (that is: $K = 12$).[14] The figures that are to follow display the median, as well as the 16 and 84 percent quantiles for the sample of impulse responses. Unless stated otherwise, these figures were constructed by generating 100,000 candidate shocks.

4.1 A somewhat more agnostic identification procedure

When we include the CPI's of the four dollarized countries in the analysis, we can first try to identify monetary US shocks by placing restrictions on the federal funds rate and *on prices in the dollarized countries only*. In particular, we can define a contractionary US monetary policy shock as a shock that does not lead to an increase in the price level in all dollarized countries, and does not lead to a decrease in the federal funds rate.

Note that this identification strategy allows us to stay agnostic with respect to the response of both US output *and prices*. Compared to Uhlig's (2005) "agnostic procedure", we thus not only leave the response of US GDP unrestricted, but also that of the US GDP deflator (which is why we refer to it as "a somewhat more agnostic identification procedure"). In this case, we thus do not avoid the price puzzle in the US by construction and allow prices to go up after a monetary contraction, if the data want this to be. We still have to restrict prices somewhere, but this is what we use the dollarized countries for. To allow for a role of the cost channel over there, we only restrict the sign of the price responses in these countries as of the fourth month after the shock has hit.

However, Figures 1 and 2 suggest that most shocks satisfying the restrictions tend to decrease prices as of impact already. In particular, Figure 2 shows that only a minor share of the identified contractionary shocks increase the US price level. We thus do not find much evidence for a significant role of the working capital channel in the US. Here, we stress that this is not imposed by construction: under this specification the US price level is left unrestricted at all horizons, but nevertheless nearly all impulse responses indicate that prices fall after the shocks we have identified. For Ecuador and El Salvador on the other hand, the cost channel seems to be somewhat more important, as there are more "contractionary monetary policy shocks" that increase their price levels on impact. This is consistent with the view that the working capital channel is more important for emerging economies than it

[14] This implies that the restrictions on prices are only imposed from the fourth month after the shock has hit up to and including the twelfth month.

is for the US, as short-term bank financing tends to play a bigger role in the former (*cf.* Van Wijnbergen (1982, p. 134)).

Second, we find relatively little inertia in the price responses, which suggests that prices were pretty flexible over the sample period: in all countries, the price response reaches its maximum (in absolute value) at very short horizons. Comparing the sizes of these responses suggests that prices in the dollarized economies were more flexible than those in the US. This is in line with the micro-level evidence presented in Morandé and Tejada (2008), who report that emerging economies typically exhibit less price rigidities than the US.

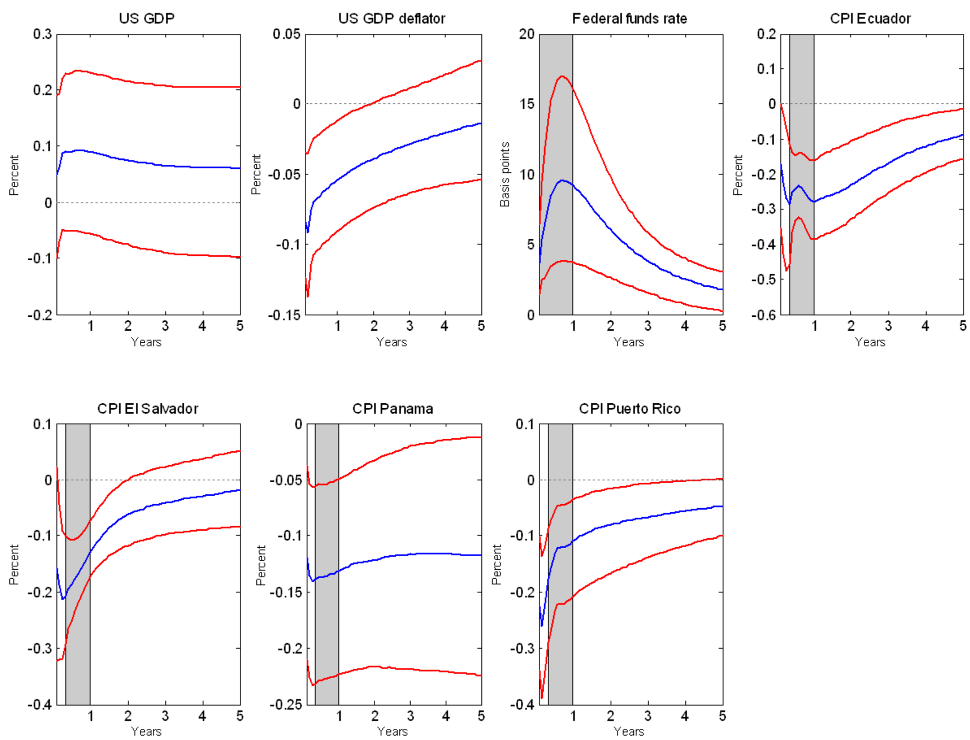

Figure 1: Impulse responses to those shocks that decrease prices in all dollarized countries, while increasing the federal funds rate. The shaded bars indicate the variables and periods where the sign restrictions are imposed.

Finally, we fail to find a clear effect of a monetary contraction on US output: there are still quite a few "contractionary US monetary policy shocks" that do not depress US real GDP. This already shows from Figure 1 (where a zero response is contained in the 16 to 84 percent quantile band), but can also be seen by looking at the distribution of all impact impulse responses for US GDP (Figure 2), which is almost symmetric around zero. We thus cannot reject monetary neutrality, which is consistent with our finding that prices seem to have been rather flexible over the sample period.

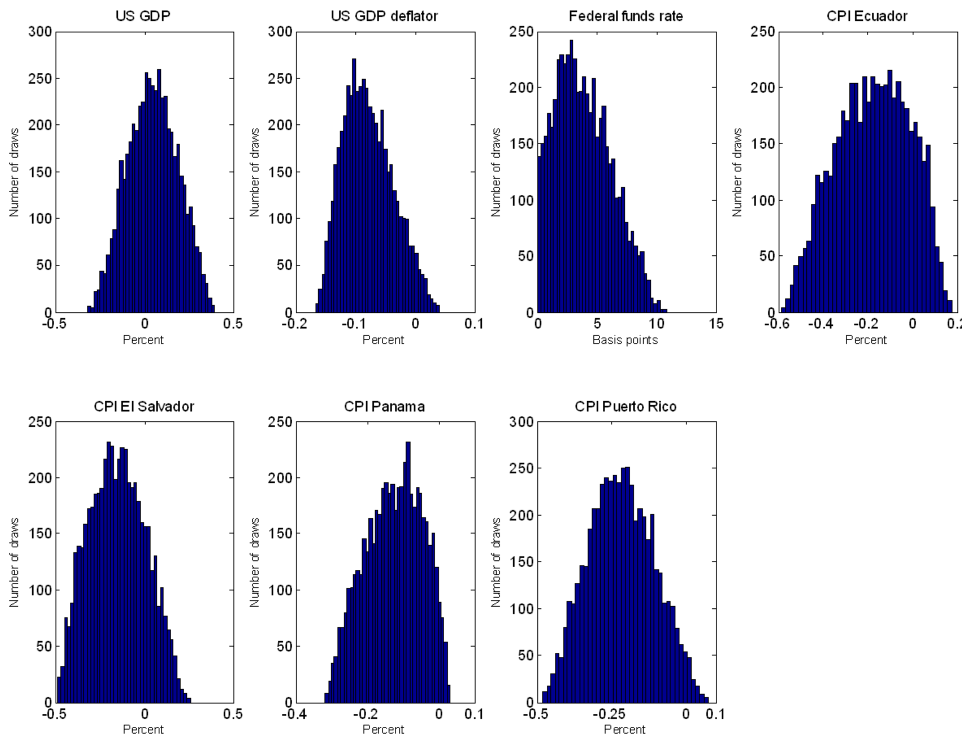

Figure 2: Distribution of the impact impulse responses underlying Figure 1.

4.2 An agnostic identification procedure

One can also increase the number of restrictions by restricting the response of the US price level as well. This specification is thus essentially the exercise in Uhlig (2005), augmented with price levels of the four dollarized countries. In this case, we define a contractionary US monetary policy shock as a shock that does not lead to an increase in prices in all dollarized countries and the US, and does not lead to a decrease in the federal funds rate.[15]

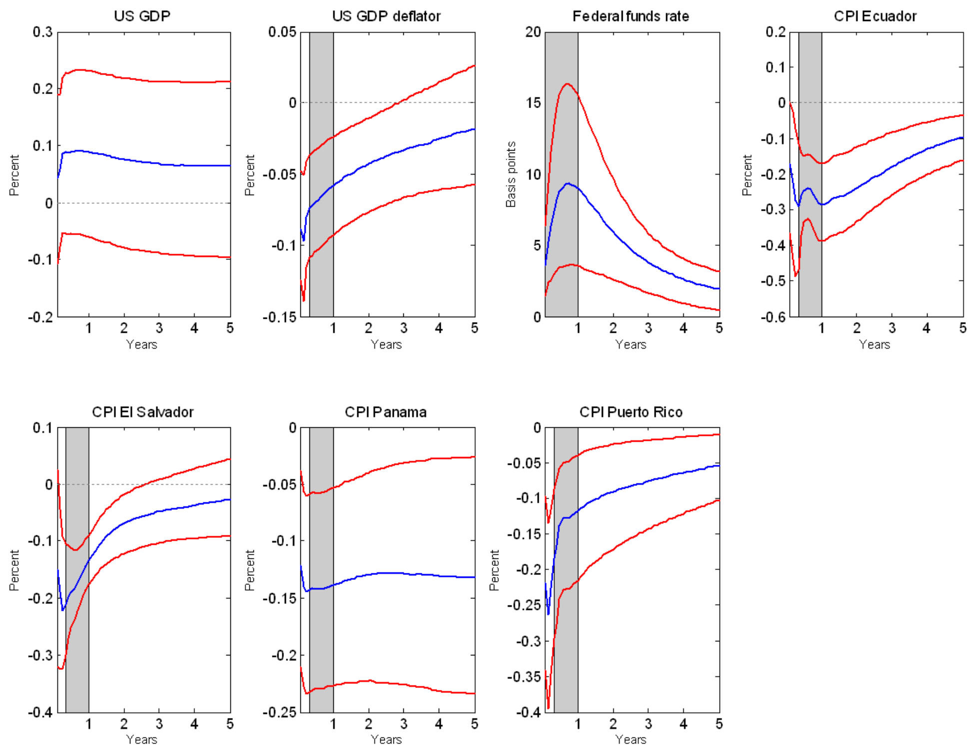

Figure 3: Impulse responses to those shocks that decrease prices in all dollarized countries and the US, while increasing the federal funds rate. The shaded bars indicate the variables and periods where the sign restrictions are imposed.

As Figure 3 shows, the additional restriction hardly affects our previous results. Apparently, nearly all shocks that increase the federal funds rate and decrease prices in the dollarized countries, also decrease the US price level. Hence, the additional restriction placed on US prices does not seem to add much information. This is consistent with our underlying idea that a true monetary policy shock should decrease

[15] As in the previous section, we only impose the restrictions on prices after four months, to allow for a short price increase after a monetary contraction as predicted by models with a working capital channel.

the price level in all countries that use the US dollar. Again, output fails to show a clear response.

As can be seen by comparing Figure 4 with Figure 2, the distributions of impact impulse responses are also hardly affected by the additional restriction.

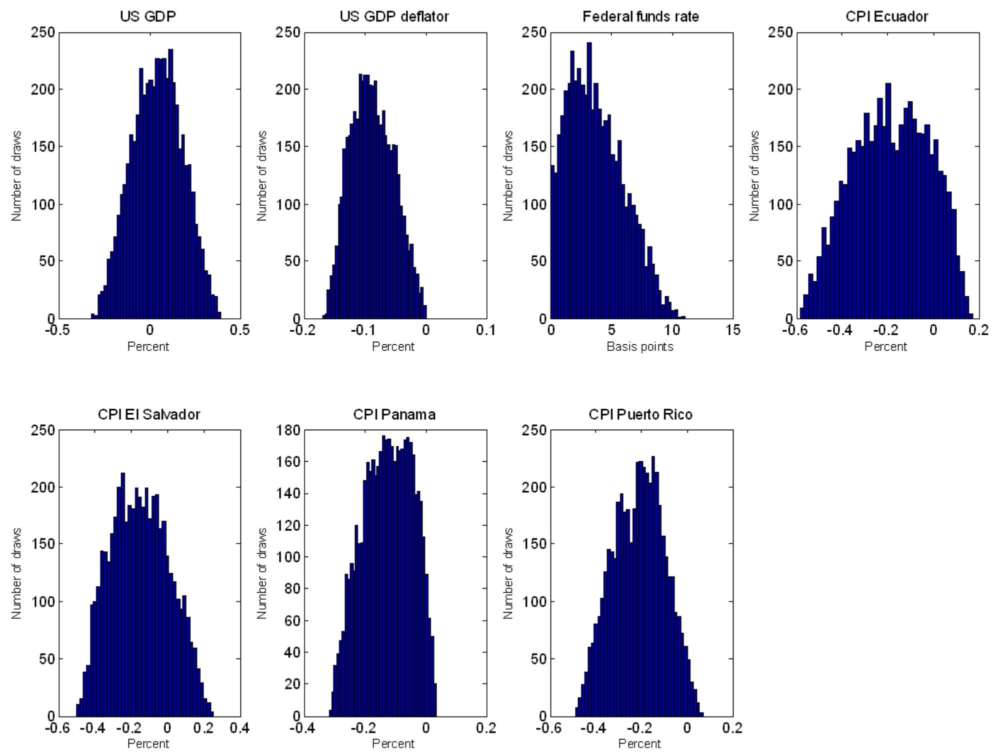

Figure 4: Distribution of the impact impulse responses underlying Figure 3.

4.3 A not-so-agnostic identification procedure

Finally, we restrict the sign of the response of the final unrestricted variable (US real GDP) and define a contractionary US monetary policy shock as a shock that does not lead to an increase in US prices and output, does not increase prices in all four dollarized countries, and does not lead to a decrease in the federal funds rate. This definition of a contractionary monetary policy shock is consistent with a standard

New Keynesian model.[16]

Of course, now the approach is no longer agnostic, but it is still interesting to see what this exercise produces. In particular, it gives an answer to the question: "How difficult is it, to find shocks that satisfy the standard New Keynesian definition of a monetary policy shock?"

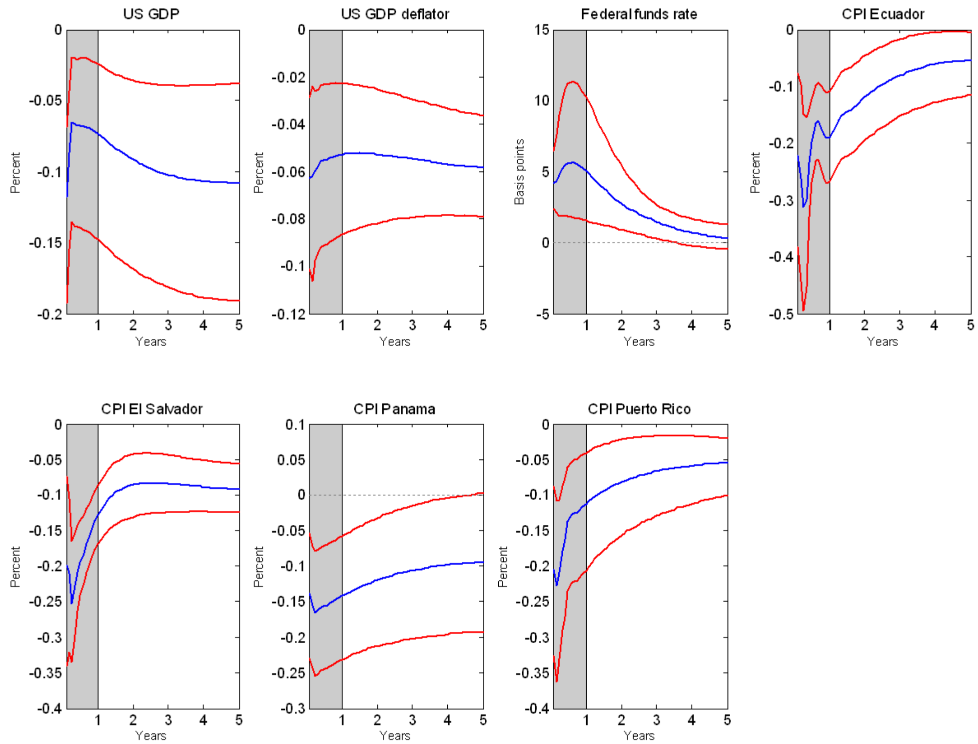

Figure 5: Impulse responses to those shocks that decrease prices in all dollarized countries, prices and output in the US, while increasing the federal funds rate. The shaded bars indicate the variables and periods where the sign restrictions are imposed.

We have found it to be pretty difficult. In this case, only 0.81 percent of all candidate shocks satisfied the imposed restrictions (against 4.77 percent in Section

[16] Here, we strictly follow the canonical New Keynesian model (which does not feature a working capital channel) and impose the restrictions on prices as of impact already. Again, the results are almost identical to those obtained when we remain slightly more agnostic by leaving prices initially unrestricted for some months.

4.2 and 4.90 percent in Section 4.1).[17] Hence, there do not seem to be a lot of shocks in the data that look like the standard New Keynesian monetary policy shocks that we encounter so often in our models.

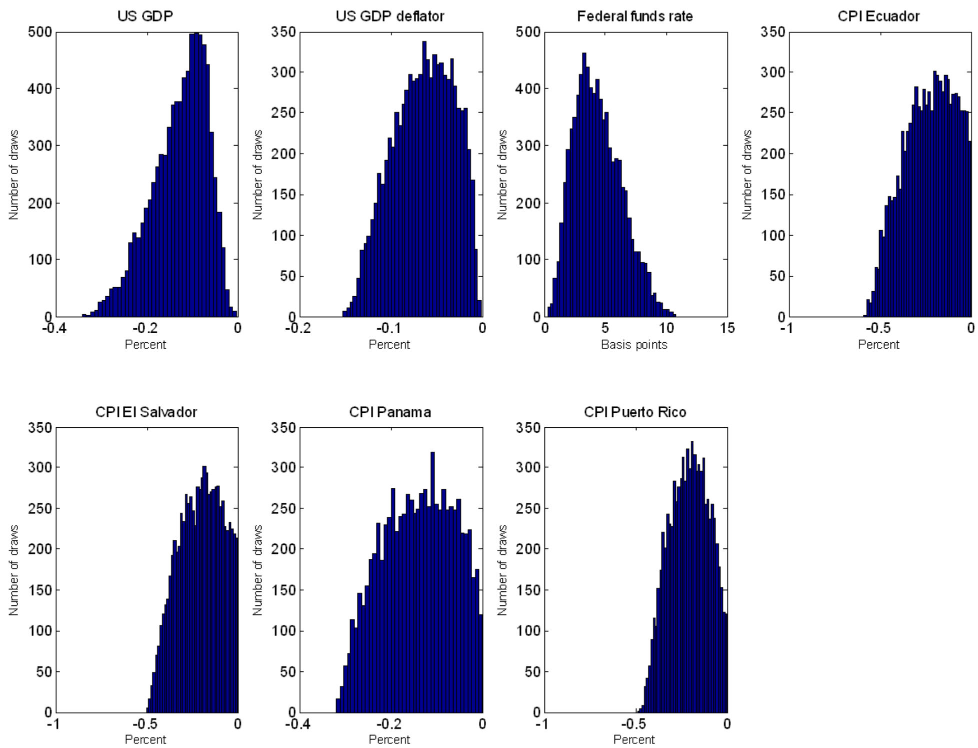

Figure 6: Distribution of the impact impulse responses underlying Figure 5.

5 Robustness

We have found the results reported in the previous section to be very robust to other sample periods, different VAR specifications and to changes in the duration of the imposed restrictions. In this section, we will describe some of the robustness checks that we have conducted. As most of the graphs resulting from the checks are very similar to the original ones, we do not include the figures in the main text, but rather

[17] Given the small acceptance rate of the candidate shocks in this case, we increased the number of candidate draws for M from 100,000 to 1,000,000.

refer to an Appendix. Unless stated otherwise, results that are to follow apply to the somewhat more agnostic identification procedure, in which the restrictions are imposed for one year after the shock (so $K = 12$). Again, when we restrict prices, we only do this as of the fourth month after the shock has hit (unless stated otherwise) to allow for the working capital view of monetary shocks.

5.1 VAR specification

First, we investigated the robustness of our results to different specifications of the VAR. We for example reproduced Figures 1 and 2 when we include a linear time trend. As one can see in the Appendix, this hardly affects the results. Similarly, one can verify that following Giordani's (2004) suggestion by estimating the VAR on the output gap and inflation (rather than on their levels, as is done in the benchmark specification) yields a very similar picture. Finally, we also checked that our results are robust to the inclusion of either a monetary aggregate like M2 (as this is argued for by Leeper and Roush (2003) and Favara and Giordani (2009) among others) or to the inclusion of nonborrowed reserves (which was called for by Christiano, Eichenbaum and Evans (1996)).[18]

5.2 Sample

As set out in Section 3, the benchmark results were generated by estimating the VAR-coefficients on samples of different lengths, so as to use all available information for each country. In the Appendix we show that the results are fully robust to limiting the analysis to the US, Panama and Puerto Rico only, in which case all coefficients can be estimated on data running from 1989:1 to 2010:7.

Moreover, the Appendix also contains results that were generated by restricting the sample to the post-2001 period, during which all countries in our study were officially dollarized. If we estimate the VAR over this period, we throw away the available data points for the US, Panama and Puerto Rico from 1989:1 to 2000:12, but as one can verify this does not seriously affect the results either.

In addition, one can also consider different start and end dates for all variables without significantly affecting the results. As reported in the Appendix, repeating the analysis on data starting in 1984:1 (just after the Volcker disinflation period), or in 1974:1 (when our series for Panama's CPI starts) does not really change anything.

[18] When we use the specification with nonborrowed reserves, we need to end the sample in 2007:12. Because of the financial crisis that followed, this variable turned negative from 2008:1 to 2008:11 (which prevents us from taking logs).

Similarly, the results are also robust to excluding the financial crisis of the late 2000s from the sample (in that case, we end the dataset in 2007:12 - when the Great Recession started according to the NBER Business Cycle Dating Committee)

5.3 Sign restrictions

Finally, we also checked the robustness of our results by varying the number of periods during which the sign restrictions are imposed (K). Uhlig (2005) finds the somewhat surprising result that the distribution of impulse responses for GDP moves up with K (so the longer the sign restrictions are imposed, the bigger the output puzzle becomes).

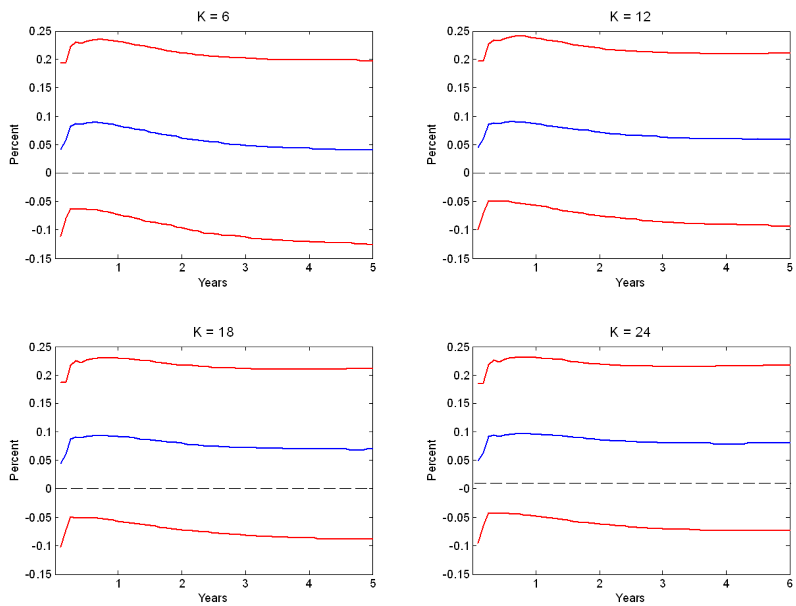

Figure 7: Impulse responses of GDP to those shocks that decrease prices in all dollarized countries, while increasing the federal funds rate for different durations of the sign restrictions.

As Figure 7 however shows, this is not the case in the present setting. Instead, the output impulse responses hardly change as we vary the duration of the restrictions. The impulse responses of the other variables are virtually unaffected by our choice of K as well (not reported).

Our results are also robust to imposing different starting periods for the sign restrictions on prices. In our benchmark setting, we only impose the restrictions on prices as of the fourth month after the shock has hit (to allow for the working capital view of monetary policy shocks). Of course, this is rather arbitrary, as economic theory only suggests that the cost channel is unlikely to play a role "at longer horizons". However, as shown in the Appendix, waiting with the imposition of the restrictions on prices until eight months after the shock has hit, or imposing them as of impact already, produces very similar results.

6 Conclusions and directions for future research

In this paper, we extend the traditional approach of identifying US monetary policy shocks through sign restrictions on US variables with additional information. In particular, we incorporate the fact that US monetary policy also affects variables in dollarized countries. This paper therefore adds price data from the latter to the analysis. This makes it possible to place sign restrictions on more variables (which should help to identify the shocks more precisely), while it is also useful in distinguishing monetary from non-monetary US shocks, as the former are transmitted rapidly to the client countries through financial markets, while the latter typically need more time to arrive in a different region.

Once we include the dollarized countries in the analysis, we identify US monetary shocks by imposing sign restrictions on the federal funds rate and *on prices in the dollarized countries only*. This approach thus allows us to remain agnostic with respect to the response of both US output *and prices*. The fact that one is able to remain agnostic on the response of US prices has some value if one is sympathetic to the working capital view of monetary policy shocks (as Barth and Ramey (2001), Christiano, Eichenbaum and Evans (2005) and Ravenna and Walsh (2006) are, among others). After all, if one believes that the working capital channel is important (so that firms need to borrow funds in order to be able to pay for their production factors), the interest rate becomes a determinant of real marginal costs and prices may show a short increase after a monetary contraction. In that case, one misidentifies the true monetary shock if a short-run positive price response is ruled out by construction. The current paper is therefore more agnostic on this issue and leaves the response of US prices unrestricted, while the sign restrictions on prices in the dollarized countries are only imposed after several months.

The results of this exercise however suggest that the working capital channel does not play a big role in the US: nearly all shocks that we identify decrease US prices as of impact already. So according to this finding, the US price puzzle does not seem to be an empirical fact that is to be matched by theoretical models. It rather seems to be a consequence of monetary policy shock misidentification (as argued in Canova and Pina (2005) and Carlstrom, Fuerst and Paustian (2009)). For Ecuador and El Salvador on the other hand, there is more evidence that the working capital channel plays a role. This is consistent with the idea that the cost channel is more important for emerging economies than it is for the US, as short-term bank financing typically plays a bigger role in the former.

Finally, we find Uhlig's (2005) conclusion that monetary shocks do not seem to have a clear effect on real GDP, to be very robust.[19] So besides detecting only very few innovations that satisfy the working capital view of monetary policy shocks in the US, we do not find many shocks that are in line with a standard New Keynesian model either.

For future work, one can think of extending the current analysis with even more variables. As this reduces the parsimony of the VAR, this calls for a factor augmented VAR-approach. Amir Ahmadi and Uhlig (2009) already do this with US variables, which allows them to impose a large number of sign restrictions on the latter. However, as argued before, variables in dollarized countries are subject to the same monetary shocks as their US equivalents. Consequently, they also contain useful information on US monetary policy. Adding data from dollarized economies to a factor augmented VAR might therefore be a logical next step along this line of research.

7 References

Amir Ahmadi, P. and H. Uhlig (2009), "Measuring the dynamic effects of monetary policy shocks: A Bayesian FAVAR approach with sign restriction", mimeo, Humboldt University of Berlin.

Bagliano, F.C. and C.A. Favero (1998), "Measuring monetary policy with VAR models: An evaluation", *European Economic Review*, 6 (10), pp. 1069-1112.

Barth, M.J. and V.A. Ramey (2001), "The cost channel of monetary transmission", in: B.S. Bernanke and K. Rogoff (*eds.*), *NBER Macroeconomics Annual 2001*, Cambridge, MA: MIT Press, pp. 199-240.

Beckerman, P. (2001), "Dollarization and semi-dollarization in Ecuador", World Bank Policy Research Working Paper No. 2643.

Canova, F. (2005), "The transmission of US shocks to Latin America", *Journal of Applied Econometrics*, 20 (2), pp. 229-251.

[19]The sign restrictions study by Faust, Swanson and Wright (2004) also fails to find a clear effect of contractionary US monetary policy shocks on US GDP (despite the fact that they are less agnostic and restrict the output response to be non-positive on impact).

Canova, F. and G. de Nicoló (2002), "Monetary disturbances matter for business cycle fluctuations in the G7", *Journal of Monetary Economics*, 49 (6), pp. 1131-1159.

Canova, F. and J. Pina (2005), "What VAR tell us about DSGE models", in: C. Diebolt and C. Krystou (eds.), New Trends in Macroeconomics, New York, NY: Springer-Verlag.

Carlstrom, C.T., T.S. Fuerst and M. Paustian (2009), "Monetary policy shocks, Choleski identification, and DNK models", Journal of Monetary Economics, 56 (7), pp. 1014-1021.

Chow, G.C. and A. Lin (1971), "Best linear unbiased interpolation, distribution, and extrapolation of time series by related series", *Review of Economics and Statistics*, 53 (4), pp. 372-375.

Christiano, L.J, M. Eichenbaum and C.L. Evans (1996), "The effects of monetary policy shocks: Evidence from the flow of funds", *Review of Economics and Statistics*, 78 (1), pp. 16-34.

Christiano, L.J., M. Eichenbaum and C.L. Evans (2005), "Nominal rigidities and the dynamic effects of a shock to monetary policy", *Journal of Political Economy*, 113 (1), pp. 1-45.

Christiano, L.J, M. Eichenbaum and R. Vigfusson (2003), "What happens after a technology shock?", NBER Working Paper No. 9819.

Cushman, D.O. and T. Zha (1997), "Identifying monetary policy in a small open economy under flexible exchange rates", *Journal of Monetary Economics*, 39 (3), pp. 433-448.

Dornbusch, R. (1976), "Expectations and exchange rate dynamics", *Journal of Political Economy*, 84 (6), pp. 1161-1176.

Faust, J. (1998), "The robustness of identified VAR conclusions about money", *Carnegie Rochester Series on Public Policy*, 49, pp. 207-244.

Faust, J., E.T. Swanson and J.H. Wright (2004), "Identifying VARS based on high frequency futures data", *Journal of Monetary Economics*, 51 (6), pp. 1107-1131.

Favara, G. and P. Giordani (2009), "Reconsidering the role of money for output, prices and interest rates", *Journal of Monetary Economics*, 56 (3), pp. 419-430.

Fry, R. and A.R. Pagan (2007), "Some issues in using sign restrictions for identifying structural VARs", NCER Working Paper Series 14.

Gabaix, X. (2011), "The granular origins of aggregate fluctuations", *Econometrica*, 79 (3), pp. 733-772.

Galí, J. (1999), "Technology, employment, and the business cycle: Do technology shocks explain aggregate fluctuations?", *American Economic Review*, 89 (1), pp. 249-271.

Giordani, P. (2004), "An alternative explanation of the price puzzle", *Journal of*

Monetary Economics, 51 (6), pp. 1271-1296.

Leeper, E.M. and J.E. Roush (2003), "Putting "M" back in monetary policy", *Journal of Money, Credit and Banking*, 35 (6), pp. 1217-1256.

Morandé, F. and M. Tejada (2008), "Price stickiness in emerging economies: Empirical evidence for four Latin-American countries", Universidad de Chile, Documentos de Trabajo No. 286.

Paustian, M. (2007), "Assessing sign restrictions", *The B.E. Journal of Macroeconomics*, 7 (1), Article 23.

Ravenna, F. and C.E. Walsh (2006), "Optimal monetary policy with the cost channel", *Journal of Monetary Economics*, 53 (2), pp. 199-216.

Rubio-Ramirez, J.F., D.F. Waggoner and T. Zha (2010), "Structural vector autoregressions: Theory of identification and algorithms for inference", *Review of Economic Studies*, 77 (2), pp. 665-696.

Schmidt, P. (1977), "Estimation of seemingly unrelated regressions with unequal numbers of observations", *Journal of Econometrics*, 5 (3), pp. 365-377.

Sims, C.A. (1992), "Interpreting the macroeconomic time series facts", *European Economic Review*, 36 (5), pp. 975-1011.

Uhlig, H. (2005), "What are the effects of monetary policy on output? Evidence from an agnostic identification procedure", *Journal of Monetary Economics*, 52 (2), pp. 381-419.

Van Wijnbergen, S. (1982), "Stagflationary effects of monetary stabilization policies: A quantitative analysis of South Korea", *Journal of Development Economics*, 10 (2), pp. 133-169.

Van Wijnbergen, S. (1983), "Interest rate management in LDCs", *Journal of Monetary Economics*, 12 (3), pp. 433-452.

Willems, T. (2011), "Using dollarized countries to analyze the effects of US monetary policy shocks", mimeo, University of Amsterdam.

Zellner, A. (1962), "An efficient method of estimating seemingly unrelated regressions and tests for aggregation bias", *Journal of the American Statistical Association*, 57 (298), pp. 348-368.