On the predictive content of the PPI on CPI inflation: the case of Mexico

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1. Introduction

It would be natural to expect that shocks to producer prices, as they spill over through the production chain, should eventually have some effect on consumer prices. This should hold true for "cost-push" shocks that are expected to appear initially during the first stages of the production chain. In this case, it would also be natural, from a statistical point of view, for producer prices to "cause" consumer prices (ie producer prices should Granger-cause consumer prices). Following these considerations, information on producer prices could therefore be useful for central banks in identifying cost-push shocks and improving forecasts of consumer prices inflation.

The international experience, however, seems to suggest that the connection between producer and consumer prices is not as close as the abovementioned rationale would imply. For example, empirical studies for the United States, such as those by Clark (1995), and Blomberg and Harris (1995), find that the producer price index (PPI) does not have a significant predictive content for the future pattern of the consumer price index (CPI). The lack of robust evidence regarding a close causal link between the PPI and the CPI, along with the fact that most central banks define their inflation targets in terms of a certain measure of consumer prices, has led some central bankers to disregard the PPI as a relevant indicator for assessing inflationary trends.

Nevertheless, there are several shortcomings in the literature concerning this issue. Among these, the most relevant are:

i) In general, the range of prices included in both producer and consumer price indices differs significantly. Indeed, it is common for PPI baskets to include mainly domestically produced goods, while CPIs include comprehensive sets of goods and services.

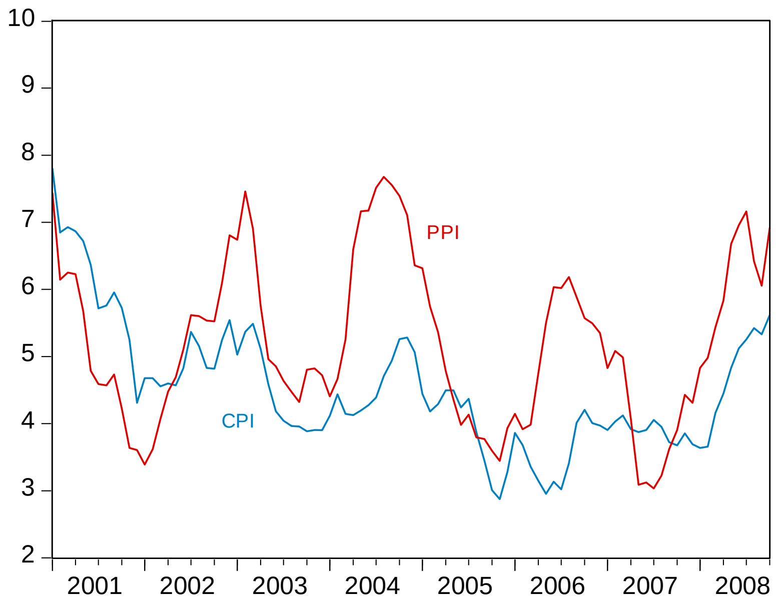
ii) The previous literature has not given enough relevance to the role played by the statistical properties and dynamic interactions of CPI and PPI time series in the analysis. In particular, most previous studies have assessed Granger-causality between these two indices by using VAR models in first differences. However, this procedure relies on two assumptions: a) price levels are I(1) series and therefore inflation rates are stationary; and, b) consumer and producer prices are not cointegrated. Should either of these two assumptions not hold, the estimation of a VAR in differences is thus not the appropriate tool for analysis. In particular, if the price-level series are I(2), then the causality analysis should take this property into account, which further complicates the study. Regarding cointegration, it is well known that, if two series are cointegrated, the VAR in first differences suffers from omitted-variable biases, because it does not include the relevant error correction mechanism (ECM) term. These biases can make Granger-causality tests lead to misleading conclusions (an issue pointed out by Granger (1988)).

[1] Banco de México. The views and conclusions expressed in this paper are solely those of the authors and do not necessarily reflect those of Banco de México. We would like to thank José Antonio Murillo and Antonio Noriega for their help and comments and Gabriel López-Moctezuma for excellent research assistance.

This note readdresses the previous evidence concerning the possibility of a causal relationship between the PPI and the CPI, using data of both price indices in Mexico. We believe this country is an appropriate case for studying the dynamic relationship between the two indices. Indeed, since 1994, the range of prices in the PPI has included the service sector, and the methodology to compute both indices is homogeneous. Although in this case the CPI and PPI still differ, the analysis should not be as affected as in other countries by issues concerning the range of prices included in such indices. For the purposes of this paper, the statistical and dynamic interactions of both time series are considered to be significantly relevant. Evidence is presented showing that from mid-2000 onwards, the inflation rates of both the CPI and PPI became stationary. The analysis is therefore restricted to the period when consumer and producer price inflation rates may be safely assumed to be I(0). The biases implicit in using a VAR in differences are explicitly avoided. We first show evidence that both PPI and CPI series seem to be indeed cointegrated and, thus, the causality analysis is based on a vector error correction model (VEC), which explicitly considers the role of the ECM term in the estimates.

Figure 1

Annual inflation: CPI vs PPI



Source: Bank of Mexico.

In contrast with previous studies, the results suggest that, in the case of Mexico, recent information on the PPI seems to be useful for improving forecasts of CPI inflation. In particular, CPI inflation responds significantly to disequilibrium errors with respect to the long-run relationship between consumer and producer prices (ie whenever producer prices suffer a shock, CPI inflation increases temporarily until consumer price levels adjust to their long-run relationship with producer prices). Thus, what may have led previous literature to conclude that the PPI is not useful in predicting CPI movements seems to be precisely the omission of this relevant transmission mechanism in the analysis.[2]

[2] The Bank of Mexico's latest experience with the PPI in assessing consumer inflationary pressures tends to confirm these conclusions. In some of the recent episodes in which the trajectory of CPI inflation has changed course, the PPI did in fact provide an early warning about the inflection point (see Figure 1).

The rest of the document is organized as follows: Section 2 analyzes the statistical properties of the CPI and the PPI series over time, and in particular, their degree of persistence. Section 3 describes the methodology used to determine the usefulness of the PPI as a predictor of CPI inflation. Section 4 summarizes the empirical results. Finally, Section 5 presents some final remarks regarding the possible lessons that may be obtained from the Mexican experience on the use of output-based price indices to assess inflationary pressures.

2. Changes in the persistence of the CPI and the PPI

In order to analyze the change in the persistence of both the CPI and PPI, the first step is to identify their basic time series properties. These properties constitute a building block for further research. It is of particular relevance to identify the order of integration of the data; that is, to assess whether PPI and CPI inflation rates are stationary I(0) processes or not. As mentioned before, if inflation rates are non-stationary I(1) processes, then the price levels would be I(2) processes, and the analysis to identify the pass-through of producer price shocks to consumer prices would therefore be more complicated.

Identifying whether inflation rates are stationary or not becomes more difficult when shifts in monetary policy, among other factors, make inflation rates switch from non-stationary to stationary regimes, or vice versa. However, several tests have recently been developed to accurately decompose the sample in a time series observed in stationary and non-stationary behavior segments. Regarding the Mexican economy, evidence based on this type of tests supports the idea that consumer price inflation shifted from a non-stationary to a stationary regime around 2000 (see Chiquiar, et al (2007)). This date nearly coincides with the period when the Bank of Mexico formally adopted an inflation targeting regime.

The latest development in this methodology is based on a test for multiple changes in persistence by Leybourne, Kim and Taylor (2007), which also allows for estimating the dates of change in a consistent way. Their test identifies all stationary periods within the sample, effectively decomposing the data into stationary (or I(0)) and non-stationary (or I(1)) subsamples. When no I(1) behavior is detected, the series is stationary. The periods identified as I(0)/I(1) can then be analyzed in terms of both timing and operating rules of monetary policy.

Table 1

Test for changes in persistence

Series	Sample	Starting date for I(0) period
CPI inflation	1994:02–2008:10	2000:05
PPI inflation	1994:02–2008:10	2000:04

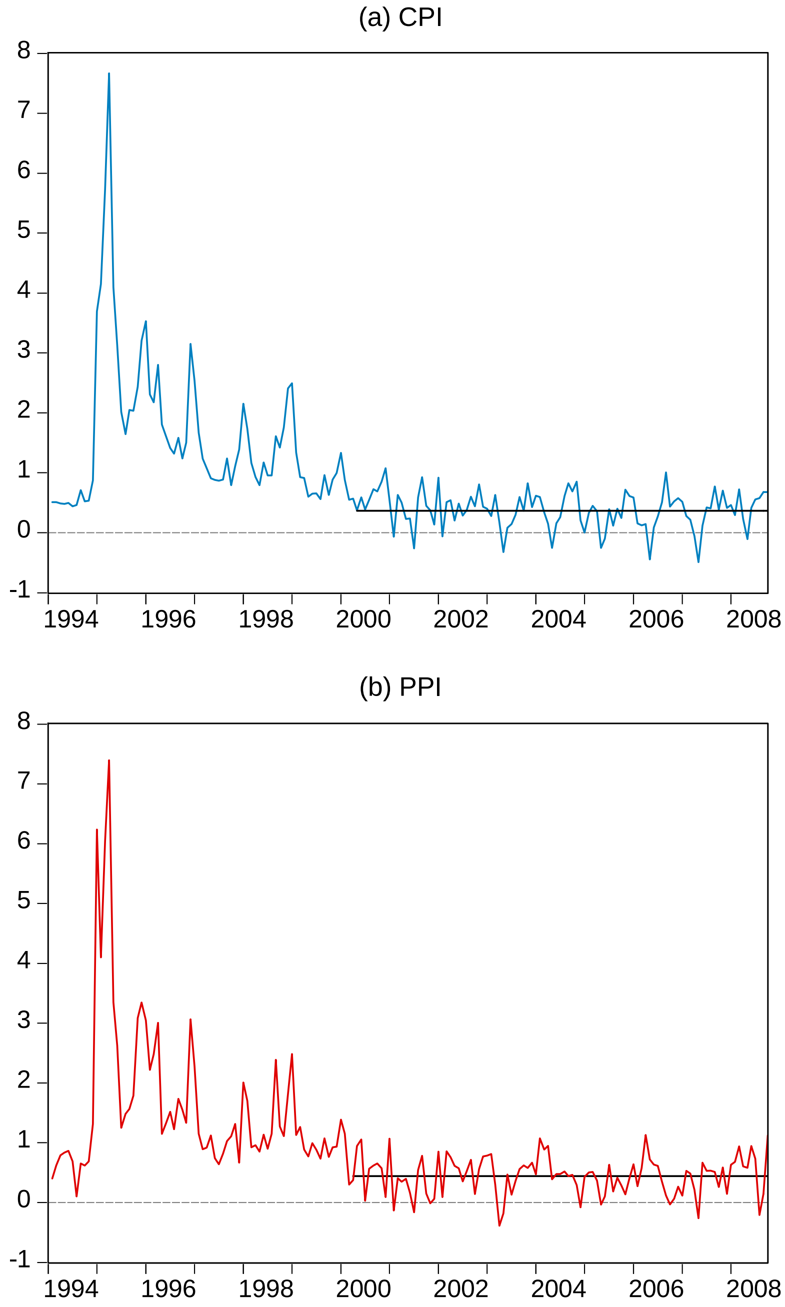
Source: Own calculations with data from Bank of Mexico.

The results of the test for monthly inflation data based on CPI and PPI inflation rates in Mexico suggest that, in both cases, inflation shifted from a non-stationary to a stationary regime around mid-2000. Table 1 summarizes the results. The second column refers to the sample to which the testing procedure was applied. The following column reports the date identified by the procedure as the beginning of the I(0) subsample. For instance, for the CPI, the test identifies a single I(0) period from May 2000. This means that from 1994:02 to 2000:04, CPI inflation seems to have behaved in a non-stationary fashion (ie as a I(1) process), while from 2000:05 onwards, the test suggests that this inflation rate behaved as a stationary process. Very similar conclusions can be reached regarding PPI inflation. Apparently, from the beginning of the sample to the year 2000, the data behaves as a non-stationary process, while from mid-2000 onwards, the inflation indices behave in a stationary way. The level of significance for all changes in persistence was 1%. These findings are similar to those reported by Chiquiar et al (2007).

Figure 2 represents the results graphically. The graphs plot the two inflation series, together with straight horizontal lines indicating the stationary period, as identified by the persistence change test. For convenience, this line is drawn at the inflation mean during the I(0) period identified by the test.

Figure 2

Monthly CPI and PPI inflation



Source: Bank of Mexico.

To conclude, the two inflation measures analyzed apparently switched from non-stationary to stationary behavior during 2000. Considering that inflation is the difference between the (log) price indices, from 2001 onwards, both price indices can be treated as I(1) variables. Given the latter, for the rest of the paper, the analysis will be conducted by restricting the sample to the period from January 2001 to the last observation available (October 2008), in order to ensure that the variables are stationary in differences (I(1) in levels) and, thus, that the conventional cointegration analysis is applicable.[3]

3. Methodology to evaluate the predictive content of the PPI for the CPI

In this paper, the methodology proposed by Granger (1969) and later popularized by Sims (1972) is used to analyze if the PPI can help forecast the CPI (ie if PPI Granger-causes CPI).

The most commonly used test of Granger causality, otherwise known in econometric textbooks and software as "Granger test", is performed under a bivariate vector autoregression (VAR), where a joint exclusion test is used. In order to investigate the predictive ability of PPI inflation for CPI inflation, the relevant equation from the VAR would be:

$$\pi_t^{CPI} = \mu_0 + \sum_{j=1}^{p} \alpha_j \pi_{t-j}^{CPI} + \sum_{j=1}^{p} \beta_j \pi_{t-j}^{PPI} + \varepsilon_t, \tag{1}$$

where ε_t is considered as white noise. The VAR is typically estimated by ordinary least squares (OLS), and the number of lags, *p,* is usually determined by using an information criterion such as the Bayesian information criterion (BIC), or Schwarz Criterion. Then, a test of the null hypothesis

$$H_0 : \beta_1 = \beta_2 = ... = \beta_p = 0, \tag{2}$$

is conducted, either with the usual F-test, or with the Wald variant.[4] If the null hypothesis is rejected, then it can be concluded that PPI inflation does Granger-cause CPI inflation. These types of tests have been used in the literature to investigate the relation between PPI and CPI inflations (eg Clark (1995)).

Engle and Granger (1987), however, show that if the variables under investigation are I(1), and a linear combination of them is I(0), that is, if the variables are cointegrated, then the series will be generated by an error-correction model. Considering the natural logarithm of the price indices ($p^{CPI} = ln($CPI$)$ and $p^{PPI} = ln($PPI$)$), their first difference will be the (monthly) inflation rate. The first equation of the VEC representation would thus be:

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3 Augmented Dickey-Fuller tests (Dickey and Fuller (1979)) for this period cannot reject the hypothesis of a unit root in each price index at the 1% level. The tests were performed using a constant and a linear trend, and the number of lags was selected using the BIC criterion, starting with 18 lags.

4 The F-test applies if ε_t is assumed to be Gaussian. However, even in such a case, the F-distribution would apply only asymptotically because the lagged dependent variables that appear as regressors make the assumption of fixed regressors untenable.

$$\pi_t^{CPI} = \mu_0 + \gamma_1(z_{t-1}) + \sum_{j=1}^{p} \alpha_j \pi_{t-j}^{CPI} + \sum_{j=1}^{p} \beta_j \pi_{t-j}^{PPI} + \eta_t, \qquad (3)$$
$$z_{t-1} = p_{t-1}^{CPI} - \varphi_0 - \varphi_1 p_{t-1}^{PPI}$$

where η_t is considered as white noise, z_{t-1} is the error correction term, and can be interpreted as the degree to which the system is out of equilibrium from the long-run relationship between the series, γ_1 is the speed of adjustment, and ϕ_1 is the cointegration coefficient. After comparing equations (1) and (3) it is clear that if the price indices are cointegrated, then equation (1) is missing the error correction term, and hence, is misspecified.

Indeed, Granger (1988) shows that a consequence of the error-correction model is that at least one of the variables in the system must be caused by z_{t-1}, (which is a function of the lagged price levels). Therefore, if two variables are cointegrated, (Granger) causation must follow at least in one direction. Granger and Lin (1995) define clearly the existence of two important sources of causation in the error-correction model (3). One originates from the effect of the error correction term (ie from the long-run relationship), if γ_1 is different from zero, and the other, from the lags of the PPI inflation rate (ie from the short-run dynamics), if βs are different from zero. Accordingly, the former is called long-run Granger causality while the latter is short-run Granger causality. If the CPI and the PPI are cointegrated, then there can be short-run causation from the PPI to the CPI, long-run causation, or both. No causation from the PPI to the CPI can also occur, although this would imply some type of causation from the CPI to the PPI.

Since the results in the previous section suggest that both price indices under study are I(1) variables in the sample since 2001, it is important to emphasize that, if the two series are shown to be cointegrated, the model in equation (1) would be misspecified if z_{t-1} is not used explicitly. In this case, if equation (1) is used, the possible relevance (in levels of significance) of the PPI as a predictor of the CPI could be missed. In extreme cases, this misspecification could invalidate completely the possible use of the PPI in forecasting the CPI (eg if both variables are cointegrated and there is only long-run causality from the PPI to the CPI).

4. Granger causality from the PPI to the CPI: empirical results

In this section, the error-correction model (3) is used to investigate the causal relation between the PPI and the CPI, in both the long and short runs. First, the series must be tested for cointegration. Once evidence of cointegration is provided, equation (3) is estimated. As a final step, significance tests on γ_1 and on βs are performed to assess causality from the PPI to the CPI. All estimations consider the period from January 2001 to October 2008, a subsample characterized by the stationarity of both CPI and PPI monthly variations (see Section 2).

To test for cointegration, we employ the methodology proposed by Engle and Granger (1987). A regression of the log CPI was run on a constant, the log PPI, and 11 (centered) seasonal dummies. Then, an augmented Dickey-Fuller test was applied to the residuals of that regression (see Table 2 for test results). The null hypothesis of a unit root in the residuals can be rejected at the 5% significance level. Thus, at the same significance level, the hypothesis that both CPI and PPI are not cointegrated is rejected.

Table 2

Cointegration test

Variables	**ADF t-stat**[a]
CPI – PPI	–3.5030**

[a] Engle-Granger (1987) test. Critical Values: 1%: –3.96, 5%: –3.41, 10%: –3.12 (following Hansen (1992)). Model with constant and 11 (centered) seasonal dummies.

Source: Own calculations with data from Bank of Mexico.

Given these results, the cointegration coefficient, ϕ_1, is then estimated using the dynamic ordinary least squares estimator proposed by Stock and Watson (1993). This is a simple procedure that produces asymptotically standard normal distributed t-values, so that inference on ϕ_1 can be performed in the usual manner.[5] The point estimate is 0.8017 with a standard error of 0.0043. Therefore, the null hypothesis that the cointegration coefficient is 0.8 cannot be rejected at the 1% level, while the hypothesis that this coefficient is 1 can be rejected at the same level. A cointegration coefficient below one implies that, in the long run, the pass-through from producer prices to consumer prices is not complete, although some considerable pass-through in equilibrium exists. This scenario could arise, for example, in a situation of monopolistic competition with non-negligible fixed costs.

Since we do not reject the hypothesis that price indices are cointegrated, it is more appropriate to estimate equation (3) rather than equation (1). The results of the estimation of the corresponding bivariate VEC (where equation (3) is the first equation of the VEC) are reported in Table 3. The cointegration coefficient is again estimated to be around 0.8. The estimates of interest correspond to equation (3), which, in the VEC reported in Table 3, corresponds to the first column, and the behavior of CPI inflation. As may be noted, the error correction term is significantly different from zero at the 5% level in the CPI inflation equation.[6] Hence, there is evidence of long-run Granger causality from the PPI to the CPI. The speed of adjustment is –0.1014, which means that a shock to the equilibrium relationship is corrected by around 10% each month, so that the total effect vanishes in less than a year. It is relevant to note that a Wald test performed on the null hypothesis that the coefficients associated with the two lagged values of PPI inflation are not significant is not rejected (the p-value is 0.8774). Therefore, no short-run (Granger) causation from the PPI to the CPI is found. This result suggests that if we had estimated a VAR in first differences without including the ECM term, we might have erroneously concluded that the PPI does not cause CPI inflation. Finally, the adjusted R-squared from this regression is slightly below 0.6, which implies that this model explains slightly less than 60% of the total variation of monthly CPI inflation.

To conclude, the results of the VEC estimates and its corresponding Granger causality tests suggest that producer prices are useful for predicting CPI inflation in Mexico. In particular, even though Granger causality tests summarized in Table 3 suggest that producer price inflation is not significant for predicting consumer price inflation in the short run, the latter responds significantly to disequilibrium errors with respect to the long-run relationship between consumer and producer prices. This means that, whenever producer prices suffer a

5 The procedure proposed by Stock and Watson is to augment the equation in levels used in the Engle-Granger tests with leads, lags, and the contemporaneous value of the difference of the (log) PPI. In this case, 4 leads and equal number of lags were used, chosen according to the BIC, from a maximum of six lags (or leads).

6 Inference in the VEC can be performed as usual given that all variables entered in the equation are stationary.

shock (ie a “cost-push” shock), consumer price inflation increases temporarily until consumer price levels adjust completely to their long-run relationship with producer prices. Indeed, as can be seen in the results summarized in Table 3, the error-correction mechanism appears significantly in the consumer price inflation equations while its coefficient in the producer price inflation equation is insignificant. This suggests that, in the long run, it is consumer prices that respond to produce price shocks, and not vice versa. In turn, this means that knowledge of shocks that affect producer prices is useful to predict future changes in consumer price inflation.

Table 3

Vector error correction estimates[a]

Sample (adjusted): 2001M04 2008M10
endogenous variables: CPI – PPI

Cointegrating equation

Cointegrating Eq:	**CointEq1**	
LCPI(–1)	1	
LPPI(–1)	–0.7944	
	[–76.2523]	
C	–1.0081	
Error correction:	**D(LCPI)**	**D(LPPI)**
CointEq1	–0.1014	0.0918
	[–2.1542]	[1.3369]
D(LCPI(–1))	0.2410	–0.5183
	[1.9134]	[–2.8220]
D(LCPI(–2))	–0.0926	0.0704
	[–0.6787]	[0.3541]
D(LPPI(–1))	0.0302	0.5271
	[0.3184]	[3.8159]
D(LPPI(–2))	–0.0468	–0.0957
	[–0.5016]	[–0.7035]
C	0.0032	0.0042
	[5.3416]	[4.7436]
Adj. R–squared	0.5732	0.1604
Granger causality (p-val)	0.8774	
Schwarz Criterion	–17.4552	

[a] t-statistics in brackets. Eleven seasonal dummies (centered) were also included in each equation.

Source: Own calculations with data from Bank of Mexico.

5. Final remarks

This note presents evidence from Mexico suggesting that the PPI may have a significant predictive content for the subsequent development of CPI inflation. The causality relation from the PPI to the CPI identified in this paper is not driven by coefficients associated with short-run dynamics, but by the long-run response of consumer prices to shocks to producer prices, which leads to a temporarily higher inflation rate until the long-run equilibrium relationship between these two indices is satisfied again. Thus, in other countries that may have price-setting characteristics similar to Mexico, finding a relevant causal relationship from the PPI to the CPI may also require the specification of a statistical model for these two series that adds a long-run cointegration relationship to the short-run dynamics of these two series. The Mexican experience described in this paper could thus be useful for other central banks seeking to uncover the dynamic relationship between producer and consumer prices.

It is relevant to mention, however, that the model estimated in this paper is fundamentally a reduced form. In particular, although we have found what seems to be a significant transmission channel from producer to consumer prices, which could potentially improve the forecasting ability of the latter, we do not claim that the model presented here is the most efficient for producing inflation forecasts. Indeed, the information concerning the development of producer prices must be combined with other relevant inflation predictors to produce efficient forecasts. What the approach taken in this paper suggests is simply that, within the full set of indicators that could be used, the PPI seems to be a valuable piece of information for assessing inflationary pressures.

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