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A Note on the Predictive Content of PPI over CPI Inflation: The Case of Mexico^{*}

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Abstract

This note studies the causal relationship that may exist between the producer price index (PPI) and the consumer price index (CPI). In contrast with previous international studies, the results suggest that, in the case of Mexico, information on the PPI seems to be useful to improve 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. These results are based on in-sample and out-of-sample tests of Granger causality, in the context of an error correction model.

Keywords: Cointegration; Forecast evaluation; Granger causality; Vector error correction.

JEL Classification: C32, C53, E31, E37.

Resumen

Esta nota estudia la relación de causalidad que podría existir entre el Índice Nacional de Precios Productor (INPP) y el Índice Nacional de Precios al Consumidor (INPC). A diferencia de lo que encuentran estudios internacionales previos, los resultados sugieren que, para el caso de México, información del INPP parece ser útil para mejorar los pronósticos de la inflación del INPC. En particular, la inflación del INPC responde significativamente a desequilibrios respecto a la relación de largo plazo entre los precios a los consumidores y el índice de precios productor. Estos resultados están basados en pruebas de causalidad a la Granger, tanto dentro de la muestra como fuera de la muestra, en el contexto de un modelo de corrección de errores.

Palabras Clave: Cointegración; Causalidad a la Granger; Evaluación de pronósticos; Vector de corrección de error.

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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. As a consequence, one should expect producer prices to “cause” consumer prices from a statistical perspective (i.e., 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). For Canada, Dion (1999) studies several core inflation indicators and finds some evidence that the industrial PPI for electrical products “contain signaling information that might be useful for monitoring purposes” (p. 1). Yet, this evidence disappears when the paper analyzes the predictive ability of other components of the industrial PPI.¹ To our knowledge, these are the only papers that formally analyze the usefulness of the PPI to forecast CPI inflation.

The lack of robust evidence concerning 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 pressures. This argument is reinforced by a casual look at publications of 24 central banks during the years 2007-2009, including inflation reports, communiqués, and minutes, in which we find that 19 mentioned producer price indices, but that only 6 reference them as indicators of inflationary pressures (e.g., cost-push pressures

¹ Dion (1999) surveys papers that find some evidence of the relation between the PPI and the CPI in Canada, but those papers do not concentrate on the predictive content of the PPI.

or inflation in the “pipeline”).² This is telling considering that the period from 2007 to 2009 was a period of long swings in commodity prices.

Despite the lack of evidence on their usefulness and their limited use in central banking, the importance of identifying all relevant underlying pressures in the evolution of inflation indicators, along with several shortcomings in the literature, warrant revisiting this issue. Among the limitations in the literature, 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 goods, while CPI's include comprehensive sets of goods and services.
- ii) The literature has not given enough relevance to the role played by the statistical properties and dynamic interactions of the 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 would not be 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 bias, because it does not include the relevant error correction mechanism (ECM) term. This bias can make Granger-causality tests lead to false conclusions (an issue pointed out by Granger (1988)).

² The central banks include the Federal Reserve, the Bank of Japan, and 22 inflation targeters: Australia, Brazil, Canada, Czech Republic, Chile, Colombia, the European Central Bank, Hungary, Indonesia, Israel, Korea, Norway, New Zealand, Peru, Philippines, Poland, South Africa, Sweden, Switzerland, Thailand, Turkey, and The United Kingdom.

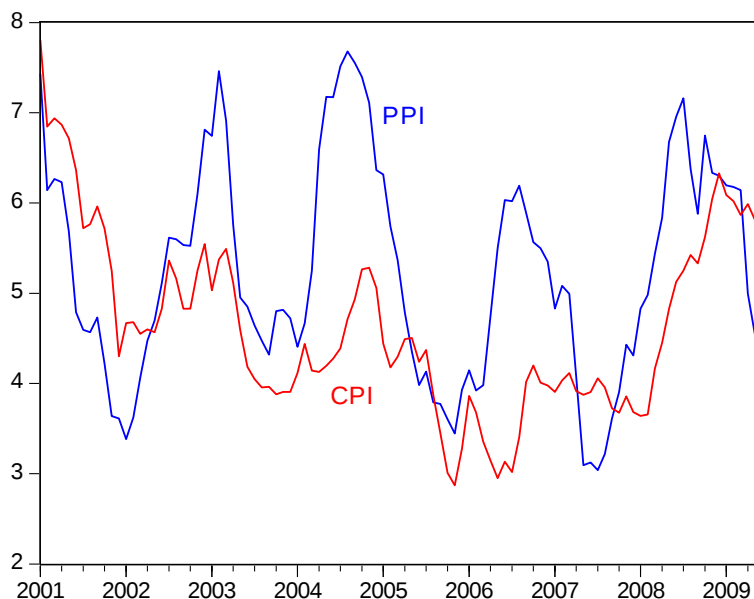
This note readdresses the previous evidence concerning the possibility of a causal relationship between the PPI and the CPI, using data of both price indices in Mexico. We believe this country is an appropriate case for studying the dynamic relationship between the two indices, as the PPI has included prices of the service sector since 1994, and the methodology to compute both indices has been homogeneous. 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 bias implicit in using a VAR in differences is 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. We present in-sample and out-of-sample evidence to support our conclusions.

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 (i.e., 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 PPI is not useful to predict CPI movements seems to be precisely the omission of this relevant transmission mechanism in the analysis. 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 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 in-sample results. Section 5 presents out-of-sample evidence. Finally, Section 6 contains 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.

Figure 1
Annual Inflation: CPI vs. PPI



Source: Bank of Mexico

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 follow a non stationary $I(1)$ process, then the price levels would follow an $I(2)$ process, 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 regimes, among other factors, make inflation rates switch from non-stationary to stationary regimes, or vice versa. However, several tests have been developed recently to accurately decompose the sample in stationary and non-stationary 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 $I(0)$ or $I(1)$ can then be analyzed in terms of both timing and operating rules of monetary policy.

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 the 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 $I(0)$ sub-sample. 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 (i.e., 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 the 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 Capistrán and Ramos-Francia (2009) and Chiquiar et. al. (2007).³

Table 1
Test for a Change in Persistence

Series	Sample	Starting date for I(0) Period
CPI inflation	1994:02- 2009:06	2000:05
PPI inflation	1994:02- 2009:06	2000:04

Source: Own calculations with data from Bank of Mexico.

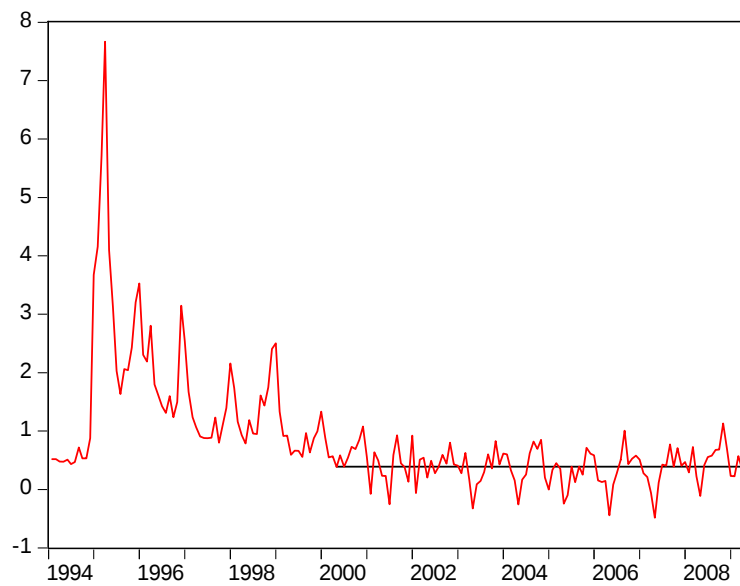
Figure 2 represents the results graphically. The graphs plot the two inflation series, together with horizontal lines indicating the stationary period, as identified by the persistence change test. For convenience, this line is drawn at the inflation mean during the I(0) period identified by the test.

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 note the analysis will be conducted by restricting the sample to the period from January 2001 to June 2009, in order to ensure that the variables are stationary in differences (I(1) in levels) and, thus, the conventional cointegration analysis is applicable.⁴

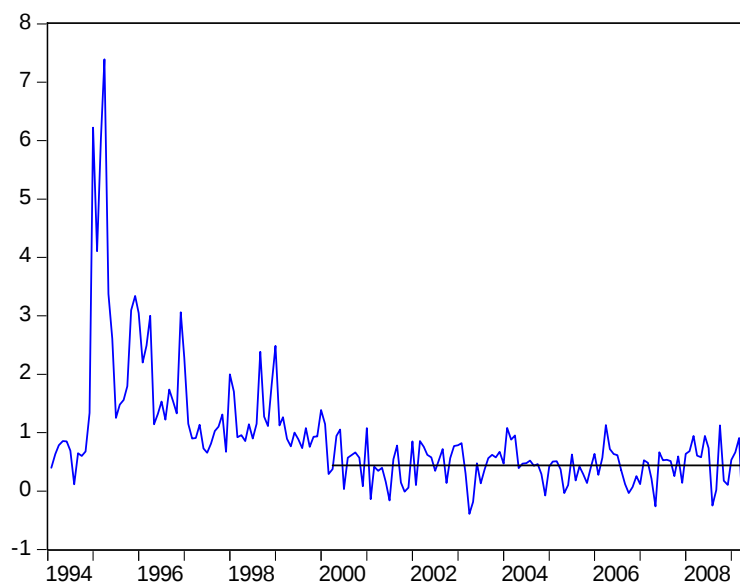
³ For evidence on changes in inflation persistence for other countries see Noriega and Ramos-Francia (2009).

⁴ 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 were selected using the BIC criterion, starting with 18 lags.

Figure 2
Monthly CPI and PPI Inflation
(a) CPI



(b) PPI



Source: Bank of Mexico

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

In this section, the methodology proposed by Granger (1969) and later popularized by Sims (1972) is used to analyze if the PPI can help forecast the CPI (i.e., 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:

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

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). Then, a test of the null hypothesis:

$$H_0 : \beta_1 = \beta_2 = \dots = \beta_p = 0, \quad (2)$$

is conducted, either with the usual F-test, or with the Wald variant.⁵ If the null hypothesis is rejected, then it can be concluded that PPI inflation does Granger-cause CPI inflation. These type of tests have been used in the literature to investigate the relation between PPI and CPI inflations (e.g., 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

⁵ 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.

the price indices ($p^{CPI} = \ln(\text{CPI})$ and $p^{PPI} = \ln(\text{PPI})$), their first difference will be the (monthly) inflation rate. The first equation of the VEC representation would thus be:

$$\begin{aligned}\pi_t^{CPI} &= \mu_0 + \gamma_1(z_{t-1}) + \sum_{j=1}^p \alpha_j \pi_{t-j}^{CPI} + \sum_{j=1}^q \beta_j \pi_{t-j}^{PPI} + \eta_t, \\ z_{t-1} &= p_{t-1}^{CPI} - \phi_0 - \phi_1 p_{t-1}^{PPI}\end{aligned}\tag{3}$$

where η_t is considered as white noise, z_{t-1} is the error correction term, which 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 (i.e., from the long-run relationship) if γ_1 is different from zero, and the other, from the lags of the PPI inflation rate (i.e., from the short-run dynamics), if β s 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 PPI to CPI, long-run causation, or both. No causation from the PPI to the CPI can also occur, although this would imply at least long-run causality from CPI to 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, if both variables are cointegrated and there is only long-run causality from the PPI to the CPI, this misspecification could lead a researcher to conclude that the PPI is useless to forecast the CPI, when in fact it is useful.

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 β_s are performed to assess causality from the PPI to the CPI. All estimations consider the period from June 2000 to June 2009, 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 with 1 lag, selected according to BIC from a maximum of 3 lags, was applied to the residuals of that regression (see table 2 for test results). The null hypothesis that both CPI and PPI are not cointegrated is rejected at the 10% significance level.

Table 2
Cointegration test

Variables	ADF t-stat ^{a/}
CPI - PPI	-3.3391*

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.⁶ The point estimate is 0.8196 with a standard error of 0.0047. With these estimates, the null hypothesis that the cointegration coefficient is 1 can be rejected at the 1% 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 exists in equilibrium. 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 with the number of lags selected using the BIC from a maximum of 3.⁷ We immediately note that the cointegration coefficient is again estimated to be around 0.8. The estimates of interest correspond to equation (3) above, 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.⁸ Hence, there is evidence of long-run Granger causality going from the PPI to the CPI. The speed of adjustment is -0.0691, which means that a shock to the equilibrium relationship is corrected by around 7% each month, so that the total effect vanishes in about a year. We do not find short-run (Granger) causation from PPI to CPI, as can be seen from the result of the t-test associated with the first lag of PPI inflation in the equation for CPI (p-value is 0.7387). 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 PPI inflation. Finally, the adjusted R-squared from this

⁶ 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, 3 leads were used and equal number of lags, chosen according to BIC, from a maximum of 3 lags (or leads).

⁷ The estimation was carried out following Johansen (1995).

⁸ Inference in the VEC can be performed as usual given that all variables in the equation are stationary.

regression is slightly below 0.6, which implies that this model explains slightly less than 60% of the total variation of monthly CPI inflation.

Table 3
Vector Error Correction Estimates ^{a/}

Sample (adjusted): 2000M10 2009M06		
Endogenous Variables: CPI - PPI		
Cointegrating Equation		
Cointegrating Eq:	CointEq1	
LCPI(-1)	1	
LPPI(-1)	-0.8203	
	[-46.6313]	
C	-0.8873	
Error Correction:	D(LCPI)	D(LPPI)
CointEq1	-0.0691	0.0336
	[-2.2793]	[0.7067]
D(LCPI(-1))	0.1962	-0.3147
	[1.8727]	[-1.9156]
D(LPPI(-1))	0.0452	0.2727
	[0.6395]	[2.4599]
C	0.0029	0.0043
	[6.2179]	[5.9508]
Adj. R-squared	0.6443	0.1968
Schwarz Criterion	-17.2298	

a/ t-statistics in brackets. 11 seasonal dummies (centered) were also included in each equation.

Source: Own calculations with data from Bank of Mexico.

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 shock (i.e., 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 producer 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.

5. Granger causality from the PPI to the CPI: out-of-sample evidence

While the regression test shows that lagged PPI help explain movements in CPI, if the production chain links consumer prices to producer prices, then producer prices should be useful for forecasting consumer prices out of sample (Clark (1995)). Given the in-sample evidence presented above, that indicates a relation between PPI and CPI may exist in the long-run, producer prices should help predict consumer prices in long-horizons. In this section we do out-of-sample Granger causality tests in order to provide additional evidence on the linkage between producer and consumer prices.

In-sample Granger tests such as the one presented before have at least two possible shortcomings when compared to the original idea of Granger causality. The first is that they are in-sample tests, whereas Granger causality is a forecasting concept that should arguably be tested out of sample. The second is that the forecasting horizon, h , is typically restricted to be one-step-ahead. In order to overcome these shortcomings, we do a test of Granger causality that is out of sample and involves multiple forecasting horizons.

In order to assess the marginal predictive power of the PPI for the CPI we forecast the h -month ahead annualized change of the CPI index,

$$\pi_{t+h}^{CPI,h} = \left[\log(p_{t+h}^{CPI}) - \log(p_t^{CPI}) \right] \frac{100}{h},$$

using two different models. First, we estimate the model:

$$\pi_{t+h}^{CPI} = \mu_1 + \sum_{j=0}^p \delta_j \pi_{t-j}^{CPI} + v_{t+h}, \quad (\text{Model 1})$$

where v_{t+h} is considered a white noise process. This model is a simple autoregressive model for CPI inflation, and has proved to be a good benchmark model to forecast this inflation (see for example Capistrán et. al. (2009)). Second, we estimate an autoregressive model augmented with data from the PPI:

$$\pi_{t+h}^{CPI} = \mu_2 + \sum_{j=0}^p \phi_j \pi_{t-j}^{CPI} + \sum_{j=0}^q \gamma_j \pi_{t-j}^{PPI} + \lambda_1 p_t^{CPI} + \lambda_2 p_t^{PPI} + \xi_{t+h}, \quad (\text{Model 2})$$

where ξ_{t+h} is a white noise process. The current levels of the CPI and the PPI (in logs) are included to take into account the error-correction term from equation (3). Notice that both models employ a direct approach to multi-step forecasting, that is, we are using horizon-specific linear models in which the dependent variable is the multi-step-ahead variable of interest.

The models are estimated by OLS, using rolling samples.⁹ Then, we generate h -step-ahead forecasts for a period that was, on purpose, left aside for evaluation. Finally, we compute root mean squared forecast errors (RMSFE) for each model and forecasting horizon.

In this context, the out-of-sample null hypothesis of Granger-non-causality is akin to the null hypothesis of a predictive ability test (Diebold and Mariano (1995)):

$$H_0 : E[e(M1, h)^2] = E[e(M2, h)^2], \quad (6)$$

where $e(M1, h)$ refers to the out-of-sample forecast error made by Model 1 for horizon h . The null hypothesis corresponds to no difference in predictive ability between the models, in the sense that the mean squared error is the same. Hence, under the null, information in the PPI would not be useful to forecast CPI inflation. Since model 1 is nested in model 2,

⁹ The first R observations are used to construct an initial set of estimates that are then used for the first prediction. The second prediction is obtained using a set of estimates based on a sample running from observation 2 to $R+1$, and so forth.

we use the test proposed by Giacomini and White (2006), which can readily be used to perform this type of out-of-sample Granger causality test.¹⁰ Notice that this test can also be interpreted as an out-of-sample test of strong exogeneity in the sense of Engle et. al. (1983).

The results using an out-of-sample period from June, 2003 to June, 2009 are presented in Table 4. The forecasts are for horizons of 1, 4, 8, 12, and 16 months-ahead. Results are presented in terms of the RMSFE for each model and horizon. The ratio of the RMSFE is also shown, with the quantity corresponding to Model 1 in the denominator. Furthermore, the p-value corresponding to the Giacomini-White test is also offered. A RMSFE ratio below one implies that Model 2, the model augmented with data from the PPI, has a smaller RMSFE. If it is accompanied by a small p-value, then this difference can be considered statistically significant. Table 4 contains the results using a 4 years (rolling) window in panel (a) and a 5 years (rolling) window in panel (b).¹¹

Table 4
Out of Sample Forecast Evaluation^{a/}

Sample: 2003M06 2009M09					
Panel (a): 4 year rolling window					
Horizon	1	4	8	12	16
RMSE M1	0.2181	0.2731	0.1982	0.1462	0.1395
RMSE M2	0.2432	0.3635	0.1597	0.1067	0.1315
RMSE Ratio	1.1148	1.3312	0.8057	0.7296	0.9426
GW p-value	0.2004	0.0008	0.1942	0.0283	0.8146
N Obs	73	70	66	62	58
Panel (b): 5 year rolling window					
Horizon	1	4	8	12	16
RMSE M1	0.2070	0.3176	0.1938	0.1605	0.1954
RMSE M2	0.2318	0.3587	0.1691	0.1063	0.1320
RMSE Ratio	1.1196	1.1293	0.8725	0.6619	0.6757
GW p-value	0.1665	0.2093	0.2898	0.0611	0.1798
N Obs	73	70	66	62	58

a/ 11 seasonal dummies (centered) where also included in each Model.

Source: Own calculations with data from Bank of Mexico.

¹⁰ See also McCracken (2007) for a discussion of out of sample tests of Granger causality.

¹¹ Since we have different number of observations for each horizon, the RMSFEs are comparable for a given horizon, but not across horizons.

Our results indicate that for short horizons (less than 8 months), the model augmented with PPI information is not useful to improve the forecasts of the simple autoregression, and hence we reject out-of-sample Granger causality. However, and in line with our previous (in-sample) results, Model 2 seems to deliver predictions with a smaller RMSFE for horizons above 8 months. In fact, panel (b) shows that for 12-months-ahead the improvement in RMSFE is above 30%, and it is statistically significant at 10%.¹²

We conclude that the PPI does indeed help to forecast CPI inflation, but that this happens for horizons beyond 8 months, and provided the levels of both indices are included in the forecasting equation.

6. 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 note 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 note could thus be useful for other central banks seeking to uncover the dynamic relationship between producer and consumer prices.

In contrast to what has been found for other countries, we have found what seems to be a significant transmission channel from producer to consumer prices, which appears to

¹² There is one case in which using CPI information alone is better than using information from the PPI and the difference is statistically significant: horizon 4 in panel (a).

improve the forecasting ability of the latter for long horizons. However, we do not claim that the model presented here is the most efficient to produce 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 note 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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