PRICE CONTROLS, HYPERINFLATION, AND THE INFLATION–RELATIVE PRICE VARIABILITY RELATIONSHIP

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Abstract. Chile established between 1970 and 1973 a fixed-price policy for many products eliminated at the end of 1973 following a military coup. We use these pricing policies and their subsequent removal to determine how inflation affects relative price variability under contrasting price regimes and a hyperinflationary environment. We do so using unique monthly data for 23 food products between January 1970 and May 1982. We construct different relative price variability measures and consistently find that the sensitivity of relative price variability to inflation is larger when there is price-fixing.

JEL E31, E64, E65, N16.

1 Introduction

Governments use price controls typically to maintain affordable prices for the more impoverished population on "sensible" products, such as gasoline, the rent of apartments, or food products such as bread or rice. Artificially depressed prices produce shortages as individuals increase demand, and firms reduce supply. Hence, while they seem a good idea to increase popularity, they distort relative prices.

Economic theory suggests that relative prices act as a signal for agents to optimally assign their resources (e.g., how much they work or how much they consume). It then follows that relative price variability may impede economic agents' to allocate their resources appropriately. This extensive relative price variability can, in turn, affect welfare provided that this stems from non-fundamental factors such as hyperinflationary environments that distort prices' information content[1]. Price controls, such as wage and price freezes and exchange rate pegging, usually arise in these high inflation environments and were part of heterodox price plans followed in Latin America between the 1970s and mid-1980s (Blejer and Cheasty, 1988).

*We received helpful comments from two anonymous referees. Remaining errors are our own.

[1]There are increases in RPV that may depend on changes in fundamental variables and that are not associated with a decrease in welfare. However, when the increases in RPV are explained simply by increases in inflation, there is a distortion in relative prices and it affects welfare.

Chilean history during the 70s provides a first-hand example of this sort of heterodox policies. Under President Salvador Allende's government (1970–1973), many firms were nationalized and became subject to state control. This process was also accompanied by price-fixing schemes to avoid real wage deterioration and the corresponding effect on workers' welfare, derived from unusually large and rising inflation. While successful in the first year, Allende's policies soon failed. While costs were rising due to government expansions in nominal wages, firms' real revenues fell because of price controls. Probably forward-looking firms, expecting not to reset their prices soon enough, found it optimal to make more significant adjustments in prices than those they would have made had they been permitted to set prices freely. The arrangement could have had two consequences. First, it could have generated an accelerated inflationary environment, in which prices increased by more than they would have done with no price-fixing. Second, it could have produced relative price distortions: with individual firms making different adjustments, relative prices in the economy may indeed have deviated considerably from free-market equilibrium prices. A military regime gradually liberalized prices after a military coup took place in September 1973.

In this paper, we empirically analyze the relationship between inflation and the variability of relative prices. We make three contributions to the existing literature. First, we focus on an economy with hyperinflation where the government sets prices for several (and often all) goods. We argue that when governments fix prices, inflation generates larger relative price variability because it induces firms to make larger changes when they are allowed to reset prices. To test this idea, we construct a unique data set on 23 food product monthly prices in Chile between January 1970 and May 1982, a period marked by several price controls and an hyperinflationary environment. These 23 products represent a subset of the products used to construct the Chilean Consumer Price Index (CPI) and account for near 17 percent of the basket during the period and around 50 percent of the food category. We illustrate that, although limited in size, our 23 food products actually tracked well the aggregate inflation behavior, thus mitigating concerns about its representativeness.

Second, Chile's economic policy history allows us to explicitly evaluate whether the relationship between relative price variability and inflation is different under contrasting price regimes. The government fixed the price of the 23 products that we collected between January 1970 and October 1973. In November 1973, half of these goods' prices were liberalized, while the other half had a fixed price until December 1976. All prices in our sample were liberalized in January 1977. We exploit these liberalization policies to study the differential impact of inflation on the relative price variability of goods with a fixed price relative to those determined by market conditions. This second point represents our major contribution. We believe that this is the first work that contributes to this regard. Our first contribution, where we study the pass-through from inflation to relative price variability during hyperinflationary periods, although compelling, has already been explored in the literature (Tang and Wang, 1993; Caraballo et al., 2006; Caraballo and Dabús, 2008).

Third, we analyze in detail the concept of coefficient of variation used in literature to measure relative price variability. We exploit that our data vary over time and between products and propose other indicators that complement traditional relative price variability measures. Importantly, we construct a relative price variability measure at the product-level. These product-specific relative price variability indicators allow us to compare the sensitivity of relative price variability to inflation of goods with fixed and liberalized prices.

Our results are as follows. In line with most studies, we find that inflation increases relative price variability. However, our results also suggest that this relationship is stronger when there is price-fixing. At an aggregate level, periods with substantial price-fixing show a higher sensitivity of relative price variability to inflation. At the product-level, the relative price variability of products with a fixed-price reacts more to inflation than that of products with a liberalized price. Therefore, price-fixing schemes may be vital in explaining part (and sometimes all) of the positive relationship found in the literature that analyzes this relationship during high inflation environments.

The rest of this paper is structured as follows. Section 2 discusses the related literature. In Section 3, we briefly summarize Chile's price-fixing experience. Section 4 presents the data and discusses the evidence emerging from these data. In Section 5, we explain our empirical strategy and discuss our results. Finally, Section 6 concludes the paper.

2 Related Literature

Several works have explored the relationship between inflation and relative price variability during the last few decades, yet there is no consensus on its actual effect. At a purely theoretical level, there are at least three mechanisms through which inflation may affect relative price variability (Aksoy et al., 2013). In menu cost models, firms face a reduction in their real revenues whenever the aggregate price is increasing. To preserve their optimum real price, they need to adjust their nominal price. However, firms cannot immediately do so, because they face a cost whenever they change a price. Furthermore, these costs vary among firms in all likelihood, which means such price adjustments create relative price distortions (Sheshinski and Weiss, 1977; Rotemberg, 1983; Alvarez et al., 2019).

A second argument stems from signal extraction models, where agents face informational problems confusing changes in absolute and relative prices (Barro, 1976; Cukierman and Wachtel, 1982). For example, an aggregate demand shock would be internalized differently across agents, producing relative price variability.[2] Finally, monetary search models point to an inconclusive link between inflation and relative price variability because of two opposing channels (Peterson and Shi, 2004; Becker and Nautz, 2012). In both cases, there exist informational asymmetries between buyers and sellers. Higher expected inflation reduces fiat money value in the first channel, increasing sellers' market power and relative price variability because buyers have incomplete information about sellers' prices. In the second channel, a given level of inflation reduces buyers' real search cost and relative price variability.

The empirical literature is also inconclusive about the real impact of inflation on relative price variability. In one of the first works on this topic, Parks (1978) argued that inflation positively affects relative price variability and that its unexpected component exerts a more considerable impact on it than its anticipated component. Cukierman and Wachtel (1982) and Van-Hoomissen (1988) provide further evidence on this effect. In a subsequent paper, Stockton (1988) showed that these works were incomplete because causality could run in both directions, and the relationship between the variables is essentially non-linear. In the same vein, Hartman (1991) argued that

[2] Although the logic of these models is different, observationally, they are almost equivalent. See Bakhshi (2002).

several restrictions were needed to produce these previous empirical findings and, thus, earlier works conveyed little to no financial information.

Notwithstanding the latter critiques, subsequent research finds that the relationship between relative price variability and inflation in mostly positive (Parsley, 1996; Fielding and Mizen, 2000; Küçük-Tuğer and Tuğer, 2004; Banerjee et al., 2007), with some exceptions that find a negative result (Silver and Ioannidis, 2001; Caglayan et al., 2008). This same research also emphasizes that this relationship may be more complicated than previously thought. The relation might be U-shaped (Bick and Nautz, 2008; Choi, 2010; Caraballo and Dabús, 2013), non- linear (Alvarez et al., 2019; Baglan et al., 2016; Rather et al., 2015), or different across countries (Aksoy et al., 2013). Also, previous models, such as New-Keynesian models, might be wrongly applied to available data (Nakamura et al., 2018).

3 Historical Context

After gaining independence in the early 19th Century, Chile maintained its relatively open, free-market economic system, with prices determined by the market, until the Great Depression.

The Great Depression hit the Chilean economy particularly hard: its exports fell by over 80 percent, and real GDP dropped by over 50 percent. Then, a period of political instability followed in the 1930s. A critical government during this period was the ",100-day socialist". This government induced the Central Bank to develop a highly expansionary monetary policy after it had suspended convertibility and gone off the Gold Standard. The "Commissary of Subsistence and Prices," Decree Law 520 (DL520)[3], was created to control the resulting price increases. This institution was in charge of fixing prices and supervise the overall price system to make necessities available to the population at "reasonable" prices.

Between 1933 and 1938, market conditions determined prices. In 1938, a center-left coalition ruled the country and adopted an import substitution development strategy that lasted until 1973. This strategy included the fixing of an undetermined number of critical prices.

In the second half of the 1960s, as part of an anti-inflationary program of the Christian Democratic government in power at the time, prices were fixed for goods and services that served as the basis for determining the consumer price index.[4]

Inheriting a slow-growing economy and an inflationary environment, the Marxist government of Salvador Allende (leading a coalition of, among others, Communists and Socialist political parties) made an effort to transform the then mixed-market economy into a centralized economy. Aggregate demand increased, and prices were fixed to achieve both product expansion and relative price stability. To prevent prices from rising with wages, the government implemented a much stronger battery of price-fixing schemes, covering more than 3000 products at one point (Wisecarver, 1986).

During his first year in office, Allende's government reduced inflation: the average inflation rate between 1970 and 1971 was around 23 percent. At the beginning of 1972, inflation began to

[3]Decree Law 520 of 1932, Socialist Government.

[4]Supreme Decree N1379, October 1966, Art. 11. Even then, the percentage changes in the price index in 1965 and 1966 were slightly higher than authorized price increases. Indeed, the average inflation rate under the government of President Frei Montalva is estimated to have been around 27 percent.

rise at elevated rates of the order of 25 to 40 percent. The second semester of that year witnessed an impressive increase in inflation: 114 percent in September, 143 percent in October, 150 percent in November, and 163 percent in December.

Allende's government ended with the military coup of September 11, 1973. On October 15, 1973, the military regime established a new price policy under Decree-Law 522. Most prices would be determined by market forces, although 33 items would still be fixed by the Directory of Trade and Industry (DIRINCO), based on cost studies.[5].

Also, there was a limited group of 18 other "informed" product prices, usually industrial products produced locally by monopolies. Based on this decree, the military regime freed almost 3000 prices at once in late 1973. Later on, prices would move from fixed to informed, and then back to free, including the prices of our sample of basic foodstuffs (see the next section). Then, to avoid moves in the other direction, from free prices to informed and fixed prices, Decree-Law N3529 was promulgated in December 1980. Almost all prices were free, determined by market conditions, except those of some public utilities. Today, Chile has one of the most open economies to international trade, with a maximum import duty of 6 percent and an average duty of less than 1 percent. It implies that the prices of tradable goods tend to be heavily influenced by international market conditions.

4 Data and Evidence

4.1 Data Description We collected monthly data for 23 food products from January 1970 to May 1982. These data come from the Statistical Yearbook ("Anuario Estadístico") and the Synthesis of the Statistical Yearbook ("Síntesis del Anuario Estadístico") of the National Institute of Statistics (Instituto Nacional de Estadísticas, INE). The products include oil, garlic, peas, onions, rice, sugar, coffee, tea, flour, eggs, milk, lettuce, butter, oranges, apples, bread, potatoes, bananas, cabbage, carrots, and three types of meat. The data consist of the current price per unit of each item in each month. These products are a small subset of the Chilean CPI basket in 1969, the one used to compute the CPI during the period that we analyze. In particular, these 23 products account for around 17 percent of the total expenditure on goods included in the basket. In other words, the sum of the expenditure weights of these 23 products adds up to 17 percent. They also account for around 50 percent of the food category. We turn to the representativeness of these 23 products in the next subsection.

We also use two macroeconomic variables available for the entire analysis period at a monthly frequency. We obtain data on national economic activity from Díaz (2006). As a measure of the Chilean economy's relevant external conditions, we use the copper price, provided by the Chilean Copper Commission (Comisión Chilena del Cobre). Finally, we measure inflation using the CPI index reported by Wagner and Díaz (2008), the official CPI index, adjusted based on data reported by Cortázar and Marshall (1980) for the 1970s.

[5]These items included bread, flour, sugar, oil, milk, coffee, tea, and some types of beef, among the products included in our sample, in addition to public utility services, gasoline, motorized vehicles, and copper, among others. See Wisecarver (1986) or Decree Law 522 for a detailed description of the items. Decree-Law 522 can be downloaded from http://www.leychile.cl/Navegar?idNorma=194744. Moreover, DIRINCO was the name that replaced the Commissary of Subsistence and Prices established in the 30s.

In our empirical analysis, we consider three different periods that cover January 1970 to May 1982. The first of them correspond to months that go from January 1970 to October 1973. In October 1966, the government signed a supreme decree that allowed it to set some prices to control inflation. In particular, a subset of prices was not allowed to increase more than 13 percent relative to December 1965 and 35 percent relative to December 1964 (Wisecarver, 1986). These prices corresponded to those included in the determination of the CPI index. This decree was in force until October 1973. Hence, in this first period, the government set the prices of all products in our sample.

The second period, which we call "partial liberalization", corresponds to November 1973 until December 1976. In this period, the military regime conducted several structural reforms in different areas, including price liberalization. Although price liberalization was a priority, it was not carried out all at once: of the 23 products in our sample, the prices of 11 were kept fixed, even after the military coup, and only later were gradually freed (January 1977)[6].

Finally, we have a third period in which market conditions set all products' prices from January 1977 to May 1982. We label this period as "full liberalization" since there was no price-fixing in our products' sample. We stop our analysis in May 1982, because as of June of that year, the government went from a fixed to a flexible exchange rate scheme, which could have affected the volatility of prices. Therefore, to avoid these economic policy changes affecting the estimates, we use data until May 1982.

The first period corresponds to a form of price-fixing because, in October 1966, the government signed a supreme decree that allowed it to set some prices to control inflation. In particular, a subset of prices was not allowed to increase more than 13 percent relative to December 1965 and 35 percent relative to December 1964 (Wisecarver, 1986). These prices corresponded to those included in the determination of the CPI index. This decree was in force until October 1973. Allende's government imposed several additional price controls, extending to those outside the realm of the CPI index, that ended abruptly with the military coup on September 11, 1973. The "partial liberalization" sample ranges from November 1973 to December 1976. In this period, the military regime conducted several structural reforms in different areas, including price liberalization. Although price liberalization was a priority, it was not carried out all at once: of the 23 products in our sample, the prices of 11 were kept fixed, even after the military coup, and only later were gradually freed (January 1977). We call the period between January 1977 until May 1982 "full liberalization" since there was no price-fixing in our sample. [7]

4.2 A First Look at the Data In this subsection, we provide raw evidence of price-fixing in our sample. We then show how these price-fixing schemes may be related to relative price variability and inflation. Thus, this subsection provides the basis for our empirical approach.

We denote $P_{i,t}$ as the nominal price of item i at time t. In addition, let P_t be the aggregate

[6] These products were flour, bread, oil, milk, butter, sugar, coffee, tea, and three types of meat. Coffee is a particular case because it was not completely fixed but "informed." We chose to include it in the list of liberalized prices. However, our main results are not affected by this choice.

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price level at time t. We measure P_t by using the aggregate price index of inflation (CPI index). Thus, we can define the relative price of product i at time t ($R_{i,t}$) as $R_{i,t} = P_{i,t}/P_t$.

Figure (1) shows the evolution of the price levels of flour and rice to understand how price fixing arises in our data. We indexed each price to its value in January 1970. For expositional purposes, the y-axis shows the log of these indexed prices. This figure shows clear evidence of price intervention in the pre-liberalization period. In particular, both flour and rice present a staggering price structure as expected in the price-fixing regime. In December 1973, the price of rice was liberalized while the price of flour remained within the group of products with a fixed price regime. Both prices increased sharply in December 1973. During 1974, the price of flour remained fixed again and was even set at lower levels than in December 1973 during 1974. From mid-1975 until the beginning of 1976, the price of flour rises quite smoothly, showing that price-fixing was in line with inflation, but already in 1976, the price level remained almost constant again. On the contrary, the price of rice showed a smooth increase throughout the partial liberalization period. In the full liberalization period, both prices exhibit smooth changes through the end of the sample.

Figure (2) provides further evidence on the consequences of price-fixing for the behavior of relative prices. Once again, we indexed each relative price to its value in January 1970 and shows the y-axis in a log-scale. This figure suggests that the setting of prices leads to a decrease in the relative price of both goods between 1970 and 1973, consistent with the evidence of Figure (1) where prices only showed discrete jumps at certain points in time. These discrete jumps are a significant source of relative price variability. During the partial liberalization period, the price of flour continued to exhibit a structure consistent with price-fixing i.e., showed discrete jumps at some points, while the price of rice showed smoother changes. In the full liberalization period, relative prices fluctuate near relatively stable levels, showing movements similar to general inflation in the medium term.

[Insert Figures 1 and 2 about here]

5 Relative Price Variability

5.1 Classic Relative Price Variability Measure The above section suggests unusual variability in relative prices during the fixed-price regime period. Any good with a fixed price showed a decrease in its relative price due to the increase in the general price level via inflation. For these goods, we observed jumps in their prices in specific months to compensate for past increases in inflation, which generate potent movements in their relative price. These movements, that do not occur in periods of flexible prices, suggest high relative price variability on fixed prices.

In this subsection, we analyze in more detail how relative price variability differs between price regimes. We start by measuring relative price variability (RPV), following many studies in the literature (see Fischer et al., 1981; Konieczny and Skrzypacz, 2005; Choi, 2010, among others) as in

$$RPV_t = \sqrt{\sum_{i=1}^{N} \omega_i (\pi_{it} - \pi_t)^2} \tag{1}$$

where $\pi_{it} = \log(p_{it}) - \log(p_{i,t-1})$ is the inflation of product i between time periods t and $t-1$, $\pi_t = \log(P_t) - \log(P_{t-1})$ is aggregate inflation, and ω_i is the weight on product i. We called this measure the traditional-RPV.

We pause here to explain how we construct the weights for each product i, ω_i. Recall from the data section that we only have information for 23 products. Hence, $N = 23$. We got information on the weights for each of these 23 products on total expenditure. Let call these original weights as $\{\tilde{\omega}_i\}_{i=1}^{23}$. By construction, these $\{\tilde{\omega}_i\}_{i=1}^{23}$ do not add up to one. To deal with this issue and because of informational constraints regarding the price of the other products that belong to the CPI basket, we reweighted each of the $\tilde{\omega}_i$ as follows

$$\omega_i = \frac{\tilde{\omega}_i}{\sum\limits_{i=1}^{23} \tilde{\omega}_i}$$

Intuitively, we are using our sample of 23 products *as if* it was the total sample of products, and we reweight each weight accordingly. By construction, the sum of ω_i now adds up to one, and we can call them "weight" in the usual sense. We acknowledge that this is a strong assumption. However, we note that our 23 products represent 7 percent of the total number of products used to compute the CPI (305) and yet they account for near 17 percent of the total CPI basket and around 40 percent of the total food category that itself represents 42 percent of the total CPI basket. We show below that this reweighting allows us to track the aggregate inflation quite well despite our reduced sample size, thus mitigating concerns about its representativeness. In particular, we construct a synthetic inflation index (π_t^s) as

$$\pi_t^s = \sum_{i=1}^{23} \omega_i \pi_{it}^a$$

where $\pi_{it}^a = \log(p_{it}) - \log(p_{i,t-12})$ represent the annual inflation of product i at time t.

To check the robustness of the traditional-RPV measure, we propose an additional index following Caraballo and Dabús (2013)

$$RPV_t^C = \frac{\sqrt{\sum\limits_{i=1}^{N} \omega_i (\pi_{it} - \pi_t)^2}}{1 + \pi_t} \tag{2}$$

This measure allows us to control the spurious increase in the RPV magnitude when the inflation rate is very high, resulting in periods of hyperinflation, such as the one in the early 1970s in Chile. Note that this definition is just a coefficient of variation because we divide the standard deviation of product-level inflation relative to aggregate inflation by the aggregate inflation during the period, as described by the term $(1 + \pi_t)$ in the denominator. At an aggregate level, the weighted-average of inflation at the product level should equals aggregate inflation. Hence, the denominator represents the mean of the product-level inflation distribution.

Figures (3) and (4) show the evolution of inflation, our synthetic inflation, and both measures of RPV respectively. We show these measures in a log-scale to ease the exposition. The first thing to note is how closely our synthetic inflation measure tracks aggregate inflation. This graphical idea is confirmed by the data since the correlation between both measures during this period is 97 percent. We now turn to the RPV measures and their relationship to inflation. We observe a positive relationship between both measures of RPV and inflation. Both RPV measures reach its maximum by the end of 1973 and coincide with the date of the first liberalization. Since then, both RPV measures and inflation have decreased, with RPV decreasing much less than inflation. At first sight, this suggests that RPV reacts more to inflation in periods of price-fixing.

In what follows, we carry out the empirical analysis to determine the impact of inflation on the variability of relative prices, following many previous studies in the literature. However, we emphasize how the rate of inflation affects RPV in different price regimes. Our empirical analysis will explain the variability of relative prices as a function of the inflation rate and some other macroeconomic variables. We use these variables as fundamentals that affect the evolution of relative prices. Specifically, we include the monthly economic activity index's growth rate and the copper price's growth rate to capture external variables' influence. We measure the growth rate at an annual frequency.

Additionally, the empirical analysis includes the annual inflation rate and the interactions between the annual inflation rate and dummy variables. The first interaction is the multiplication of the inflation rate with a dummy variable that takes the value one between November 1973 and December 1976 and is zero in other months. This variable corresponds to the differentiating effect of the inflation rate when partial price liberalization takes place. The second interaction corresponds to the multiplication of the inflation rate with a dummy variable that takes the value one between January 1977 and May 1982 and is zero in the rest of the periods. This interaction is the additional effect of the inflation rate when full price liberalization occurs.

Hence there are three parameters of interest. The first corresponds to the inflation rate parameter by itself, which measures the impact of inflation on RPV in price-fixing periods. The second is the sum of the previous coefficient with the interaction variable between November 1973 and December 1976. This sum is the impact of inflation on the RPV when there is partial price-fixing. Finally, the third parameter corresponds to the sum of the inflation rate coefficient and the interaction coefficient in January 1977 and May 1979. This parameter is the impact of inflation on RPV when there is complete liberalization of prices.

We run the following equation

$$RPV_t = \gamma_0 + \alpha_1 \pi_t + \alpha_2 D_t^A \pi_t + \alpha_3 D_t^B \pi_t + \mathbf{Z_t} \boldsymbol{\zeta} + \varepsilon_t \tag{3}$$

where RPV_t is the relative price variability of product i at time t, π_t is the inflation rate, $\mathbf{Z}_t$ corresponds to the macroeconomic variables, D_t^A is equal to one between November 1973 and December 1976, and zero otherwise and D_t^B is equal to one between January 1977 and May 1982, and zero otherwise. Besides, we include a lag of the dependent variable to capture short run dynamics. Finally, ε_t is an i.i.d error term.

Table 1 shows the results of estimating equation (3). Initially, we start by including only the inflation rate plus the lag on the RVP. We find a positive and significant coefficient, similar to the results obtained in several other studies. Column (2) also includes the interactions between the inflation rate and partial and full liberalization periods. The coefficients on both interactions are negative. The interaction on the full liberalization period is significant as long as the inflation rate coefficient remains positive (true in all our sample). This result indicates that the inflation rate affects the variability of relative prices in price-setting periods more strongly, while this effect weakens in liberalized prices. This conclusion is maintained in column (3) when we include the macroeconomic variables. Table 2 makes the same estimates as in table 1 but using the Caraballo and Dabús (2013) measure. The results are similar, so we do not comment on them.

[Insert Tables 1 and 2 about here]

Earlier literature suggests that the inflation-RPV linkage is non-linear in the inflation rate (see, for example, Nautz and Scharff, 2012). Studies using data from high-inflation countries also employ the inflation data's semi-log transformation to ameliorate non-normality concerns. We follow those studies to account for the Chilean hyperinflation period, and we provide estimates replacing the annual inflation rate by the logarithm of $(1+\pi_t)$, where π_t is the annual inflation rate.

These results appear in Table 3 where we use $\log(1+\pi_t^a)$ instead of the inflation rate. We see that the general conclusions do not change when we replace the inflation variable. The estimated coefficients are somewhat larger in magnitude. However, the story is similar; i.e., the inflation rate coefficient is significant and positive. The interactions for the periods on partial and full liberalization are negative, while the latter is significant. These results confirm the idea that there is a more significant relative price variability when there is a fixed price regime.

[Insert Table 3 about here]

Up to this point, we used the traditional definition of RPV and the measurement in Caraballo and Dabús (2013). Next, we argue that these measures occupy only part of the information available in our data. Therefore, we look for additional measures that may include more information, thus allowing us to exploit our sample's panel data features. To do this, we review the literature on the coefficient of variation matrices (see Albert and Zhang, 2010; Trickey, 2015; Van Valen, 1974, 2005)) and propose two additional measures to be used in the empirical analysis.

5.2 Measuring Relative Price Variability The literature so far has used an univariate measure of the coefficient of variation. However, we can get additional information if we take into account the nature of the data. Specifically, our data can be described by a time series of product-level inflation as in the following data matrix:

$$\Upsilon = \begin{bmatrix} \left(\pi_{t-n}^1 - \pi_{t-n}\right) & \left(\pi_{t-n}^2 - \pi_{t-n}\right) & \cdots & \cdots & \left(\pi_{t-n}^K - \pi_{t-n}\right) \\ \left(\pi_{t-(n-1)}^1 - \pi_{t-(n-1)}\right) & \left(\pi_{t-(n-1)}^2 - \pi_{t-(n-1)}\right) & \cdots & \cdots & \left(\pi_{t-(n-1)}^K - \pi_{t-(n-1)}\right) \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ \left(\pi_t^1 - \pi_t\right) & \left(\pi_t^2 - \pi_t\right) & \cdots & \cdots & \left(\pi_t^K - \pi_t\right) \end{bmatrix} \tag{4}$$

where $\left(\pi_j^i - \pi_j\right)$ is the inflation differential between product i and average inflation, π_j, in period j. We have available K products and $(n+1)$ periods of time.

Given that we have information of products over time, we construct a matrix of coefficients of variation. As in Trickey (2015), let the matrix Υ in equation (4) be a matrix of $(n+1)$

observations obtained from a k-varied distribution, with a variance-covariance matrix Ω. The population coefficient of variation matrix can be defined as in:

$$\Sigma = D_{\mu}^{-1}\Omega D_{\mu}^{-1} \tag{5}$$

where D_{μ} is a $K \times K$ diagonal matrix with i^{th} element $\frac{1+\pi_t}{\sqrt{\omega_i}}$ and ω_i being the weight on the i^{th} product. Using this notation, the matrix Σ becomes:

$$\Sigma = \begin{bmatrix} \omega_1 \sum_{j=0}^{n} \frac{\left(\pi_{t-j}^{1}-\pi_{t-j}\right)^2}{(1+\pi_t)^2} & \sqrt{\omega_1\omega_2}\sum_{j=0}^{n} \frac{\left(\pi_{t-j}^{1}-\pi_{t-j}\right)\left(\pi_{t-j}^{2}-\pi_{t-j}\right)}{(1+\pi_t)^2} & \cdots & \sqrt{\omega_1\omega_K}\sum_{j=0}^{n} \frac{\left(\pi_{t-j}^{1}-\pi_{t-j}\right)\left(\pi_{t-j}^{K}-\pi_{t-j}\right)}{(1+\pi_t)^2} \\ & \omega_2 \sum_{j=0}^{n} \frac{\left(\pi_{t-j}^{2}-\pi_{t-j}\right)^2}{(1+\pi_t)^2} & \cdots & \sqrt{\omega_2\omega_K}\sum_{j=0}^{n} \frac{\left(\pi_{t-j}^{2}-\pi_{t-j}\right)\left(\pi_{t-j}^{K}-\pi_{t-j}\right)}{(1+\pi_t)^2} \\ & & \ddots & \vdots \\ & & & \omega_K \sum_{j=0}^{n} \frac{\left(\pi_{t-j}^{K}-\pi_{t-j}\right)^2}{(1+\pi_t)^2} \end{bmatrix} \tag{6}$$

Further, to obtain an index that allow us to summarize the information of the multivariate coefficient of variation, we follow Van Valen (1974, 2005) and suggest the following measurement:

$$C_{vv} = [tr(\Sigma)]^{\frac{1}{2}} \tag{7}$$

It is interesting to note that when $n = 0$, so that there is a single period of time available, the coefficient of variation becomes $C_{vv} = \sqrt{\sum_{i=1}^{K} \omega_i \frac{\left(\pi_t^i - \pi_t\right)^2}{(1+\pi_t)^2}}$ which is our measure in equation (2) and the modified-traditional measure on the coefficient of variation proposed in Caraballo and Dabús (2013). The indices we used so far are cases of partial usage of the available information. Since our data covers more than one period, our coefficient of variation index becomes:

$$C_{vv} = \sqrt{\sum_{i=1}^{K}\sum_{j=0}^{n} \omega_i \frac{\left(\pi_{t-j}^{i} - \pi_{t-j}\right)^2}{(1+\pi_t)^2}} \tag{8}$$

The latter coefficient modifies the traditional measure by incorporating additional time-series information. From equation (8) we obtain two new measures. First, let us define the coefficient of variation within a product $WRPV^i = C_{VV}^i = \sum_{j=0}^{n} \frac{\left(\pi_{t-j}^i - \pi_{t-j}\right)^2}{(1+\pi_t)^2}$ and second, the total coefficient of variation as $C_{VV}^T = \sqrt{\sum_{i=0}^{K} \omega_i C_{VV}^i}$. The second index corresponds to the aggregation of the information on the within indexes, $WRPV^i$, and therefore it also includes time-series information.

We use these two new measures to retest how price-regimes affect inflation's pass-through to relative price variability. In the next subsection, we explore how our results change if we use those coefficients.

5.3 Total and Within Relative Price Variability The use of the within product relative price variability index ($WRPV^i$) might be particularly interesting in the experience we are analyzing. It was possible to fix prices in the Chilean economy (1970 - 1976); some products were fixed and later released while others were not fixed. These products also had different variability with probably smoother movements. We use $WRPV^i$ to focus on products that had fixed prices and that were later liberalized. These products allow us to see how relative price variability reacts to inflation in price-fixing versus liberalization periods. In particular, by focusing on product-level variation, we can use product fixed-effects. Doing so allows us to control other characteristics specific to the products that do not change over time and focus on the price regime change. Hence, it also seems important to study the variability within each product and not only between them, as suggested by the literature's traditional measures. On the other hand, the total index also considers the information within each product and between products, providing an indicator with sufficient information to look at the impact of inflation on relative price variability.

In this subsection, we implement our empirical approach using the $Total - RPV$ and the $WRPV$. We construct each series by occupying sample values and using 12 lags to build each product index. We use those lags as it provides us with enough degree of freedom to estimates our parameters[8].

We start by running the model for $Total - RPV$. This indicator is the alternative to the traditional one indicated by RPV, and therefore we use the same specification as in equation (3). We present the results using both the inflation rate and the $\log(1+\pi_t)$ in Table (4). Hence, the results are analogous to those presented in Tables (1), (2) and (3). Recall that the estimated coefficients correspond to short-term parameters whenever we include a lag of the dependent variable. As the lag of the dependent variable has an estimate of around 0.85, the long-term coefficients correspond to estimates multiplied by about 6.5 times. The short-term estimate of inflation during the price-fixing regime (α_1) is around 0.014 and is highly significant. When we use the variable $\log(1+\pi_t)$ this coefficient rises to 0.02. This latter result implies that this coefficient fluctuates in the range of 0.1 to 0.13. As in Tables (1) to (3), the estimate of the additional impact of inflation during the partial liberalization is negative, but not statistically significant. However, the impact of full liberalization is negative and significant in all reported columns. Furthermore, this effect's magnitude is large: it represents approximately 60 percent of the coefficient estimated for α_1. This last result suggests that under price-liberalization, the impact of inflation on relative price variability is 60 percent less than in the fixed-price regime.

[Insert Table 4 about here]

We now work with the data at the product level using the $WRPV$ index. As we discussed above, this allows us to control the products' specific characteristics through fixed effects. Also, note that the products can be in different situations concerning price regimes. The first of these states is a product under a price-fixing scheme during the fixed price period (before November 1973). Nevertheless, there are also other situations. We can have products with a free price regime during the partial liberalization period (between November 1973 and December 1976). Also, we have products with a fixed price regime in the latter period. Finally, there are products with a free price regime in the total liberalization period i.e., after December 1976. The impact of inflation on relative price variability might vary in these different states, and we should take into account that phenomena.

[8] We use information starting in 1969 to construct each index so that they begin in January 1970 as in the main sample. Our results are not affected by the inclusion of these additional observations.

For the ease of exposition, we start by analyzing the period 1970-1976 (price-fixing and partial liberalization) to test the idea that inflation varies across price regime states by running a regression of the form:

$$WRPV_{it} = \chi_i + \rho_1 D_t + \rho_2 D_{it} + \phi_1 \pi_t + \phi_2 D_i \pi_t + \phi_3 D_t \pi_t + \phi_4 D_{it} \pi_t + \mathbf{Y_t \Gamma} + \varepsilon_{it} \qquad (9)$$

where $WRPV_{it}$ is the within relative price variability of product i at time t, π_t is the annual inflation rate, χ_i is a product-fixed effect, D_t is equal to one after the partial liberalization, and zero otherwise, D_i is equal to one if the product was liberalized, and zero otherwise and $D_{it} = D_i \times D_t$ is the impact during the liberalization period for a good that switches from the price-fixed regime to a free-price regime. The matrix $\mathbf{Y}_t$ is a $1 \times K$ row vector of macroeconomic time-variant controls. Finally, ε_{it} is an i.i.d error term.

Our critical parameters of interest are ϕ_1 and ϕ_4. ϕ_1 provides the impact of inflation on $WRPV$ for fixed-price products before liberalization. ϕ_4 shows how liberalization alters the relationship between inflation and $WRPV$. Also, ϕ_2 corresponds to the effect of inflation on $WRPV$ for liberalized prices before liberalization. ϕ_3 measures the additional effect on fixed prices in the partial liberalization period.

Table (5) shows the estimates from equation (9). Column (1) corresponds to an equation where only the inflation rate is included by itself. Column (2) also includes the interactions' effects and allows us to estimate the coefficients ϕ_1 to ϕ_4. Column (3) additionally includes fixed effects by month and year. Columns (4) and (5) include a lag of the dependent variable, and to avoid inconsistency in the estimation (being a panel of data with lagged dependent variable), we use the technique developed in Arellano and Bond (1991). The estimates are very similar to those obtained in Table (4), when we used the $Total - RPV$ measure, with only time-series information. In this case, although we use panel data, the magnitudes are still very similar. The coefficients ϕ_1 and ϕ_4 are similar in Columns (1), (2), and (4). Columns (1) and (2) correspond to the long-term effects, while Column (4) is the short-term coefficient. Columns (3) and (5) show larger effects for ϕ_1 when we include year and month effects, but the coefficients remain significant and with the expected signs.

[Insert Table 5 about here]

Table (6) shows evidence from November 1973 to May 1982, covering partial liberalization and full price liberalization. We estimate equation (9) but in this case, the interpretation is somewhat different. The dummy variable (D_i) now corresponds to goods with fixed prices between 1973 and 1976 but that switches to a free price regime from January 1977 onwards. The liberalization period (D_t) takes the value one between January 1977 and May 1982 and zero between November 1973 and December 1976. Nevertheless, the parameter ϕ_1 corresponds to the impact of inflation under a fixed price regime, and the parameter ϕ_4 is effect of full liberalization on the relationship between inflation and $WRPV$. The results in Table (6) are similar to those reported in Table (5), in magnitude and significance, showing that the results are quite robust to changes in the periods.

[Insert Table 6 about here]

Table (7) shows the results when using a nested model, including all the data between 1970 and 1982. Column (2) shows that the estimate of ϕ_1, which is the long-term effect since we do not include the lag of the dependent variable in that Column, is close to 0.05. This coefficient is highly significant. Interestingly, ϕ_4 is negative and significant in both liberalization periods, being larger during total liberalization.

[Insert Table 7 about here]

Finally, we perform a final check on our estimates of equation 9. What we do next is to incorporate additional controls for the time effect. Specifically, we include dummies for each period. It allows us to capture any time effect common to all products and different from the treatment effect of price liberalization. This strategy comes at a cost: inflation varies only over time because we obtain inflation data from the CPI index. Therefore, in Table 8, we omit the estimate of ϕ_1 and report ϕ_4 for both 1973 and 1977. To omit the estimate of ϕ_1 does not prevent us from knowing the impact of inflation on RPV when prices are liberalized, because the parameter ϕ_4 captures that effect. We can do that since the interacted term $D_{it}\pi_t$ varies across products and time (month×year).

The results of Table 8 show that regardless of whether we estimate using separate periods or using the nested model, the parameters of interest are consistently negative and significant. It occurs in the first three columns that do not control for lags of the dependent variable and in Columns (4) to (6) where we include the lags of the dependent variable using Arellano and Bond (1991) method. Although the last three columns' estimated coefficients decrease in magnitude, as they are short-term effects, they maintain their significance and remain negative.

[Insert Table 8 about here]

To summarize, regardless of whether we use the traditional RPV (or its modification as in Caraballo and Dabús (2013)), which use data that vary only through time, or the new measures we proposed here, the conclusion is robust. Indeed, regardless of measurement, period, or estimation method, we found that price-fixing increases the sensitivity of relative price variability to inflation. Thus, by liberalizing prices, the impact of inflation on RPV decreases significantly, and in the case that we analyze here, this effect can reach almost a 60 percent decrease.

6 Conclusion

In this paper, we confirm previous results on the positive relationship between inflation and relative price variability (RPV). We take advantage of a unique monthly price data on 23 food products for the Chilean economy between January 1970 and May 1982 to analyze how the presence of price-fixing during the early 1970s and its subsequent removal after the military coup of September 11, 1973, affected this relationship. Importantly, not all prices were liberalized after 1973 and were only later released. This timing allowed us to exploit our data to provide evidence of the link between inflation and relative price variability in a hyperinflationary environment with price controls.

Our results suggest that the channel through which inflation affects RPV depends on the presence of price-fixing policies. Indeed, we analyze whether this relationship differs across pricing regimes. Interestingly, we find that the sensitivity of RPV to inflation diminishes following a price liberalization, with this effect possibly being quite significant.

Although it is not the main focus of this paper, we now provide a possible rationale for our result. We hypothesize that forward-looking firms made more significant price adjustments under a fixed-price regime than in a free-market environment because they expected to be constrained in their pricing decisions. As a result, product prices jumped when firms could update their prices (given by the government threshold). These discrete jumps contributed to increasing relative price variability because they did not coincide with each product. Since inflation was rising during this period at accelerated rates, not just due to price-fixing, it is intuitive that inflation mainly affects the RPV of fixed-price products, those that suffer the most from hyperinflation. Therefore, we highlight that whether prices are fixed or determined by market conditions is crucial for understanding how inflation impacts RPV.

Note that many governments often set prices in regulated industries, which might be a possible driver of existing literature results. Recently, some countries have applied extensive price-fixing policies, e.g., in Latin America, Argentina until the end of 2016, and Venezuela to the present, like Chile in the late 1960s and early 1970s. Our analysis suggests that these price-fixing policies might significantly increase relative price fluctuations, which distort economic resource allocations and might ultimately impact welfare.

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Fig. 1. Evolution of the nominal prices of flour and rice, 1970-1982

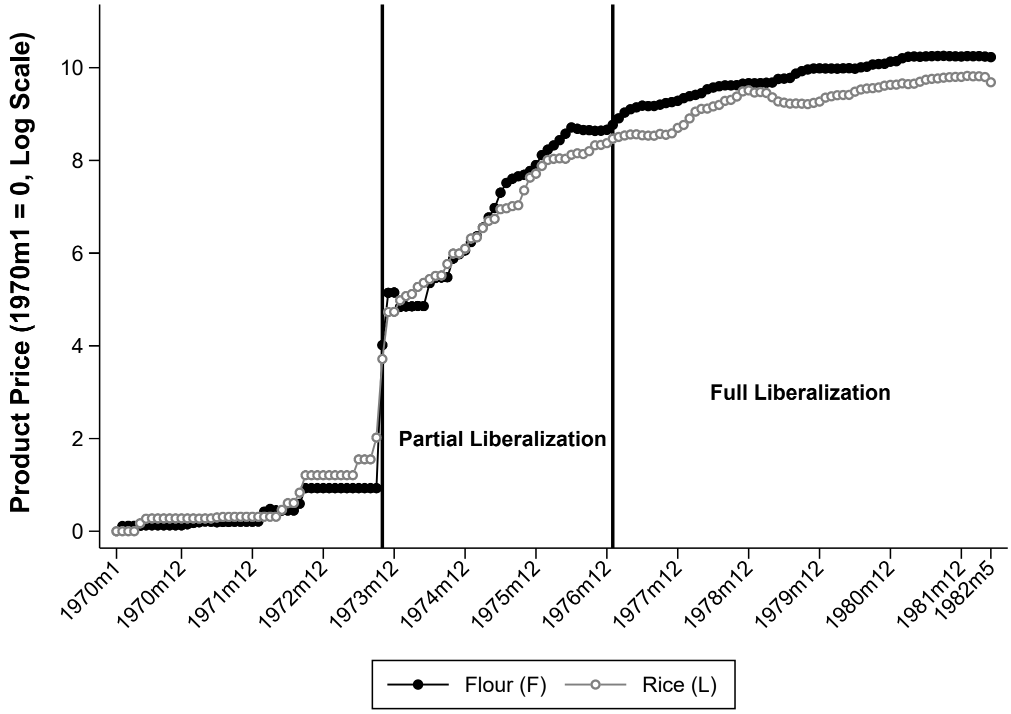

Fig. 2. Evolution of the relative prices of flour and rice, 1970-1982

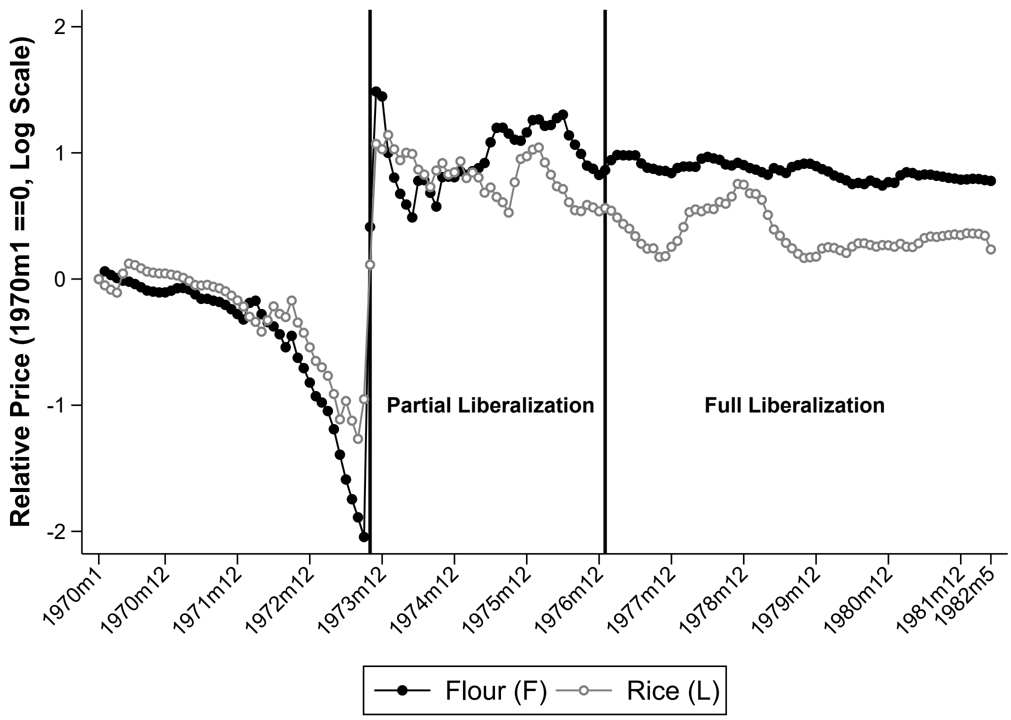

Fig. 3. Classic RPV measure

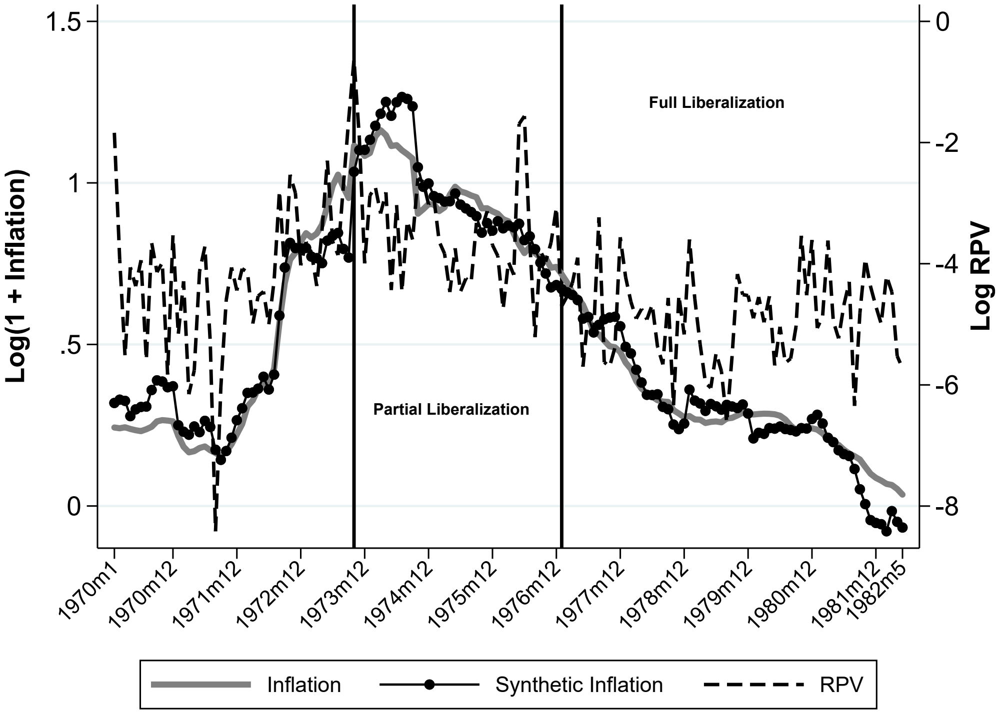

Fig. 4. Caraballo and Dabús (2013) RPV measure

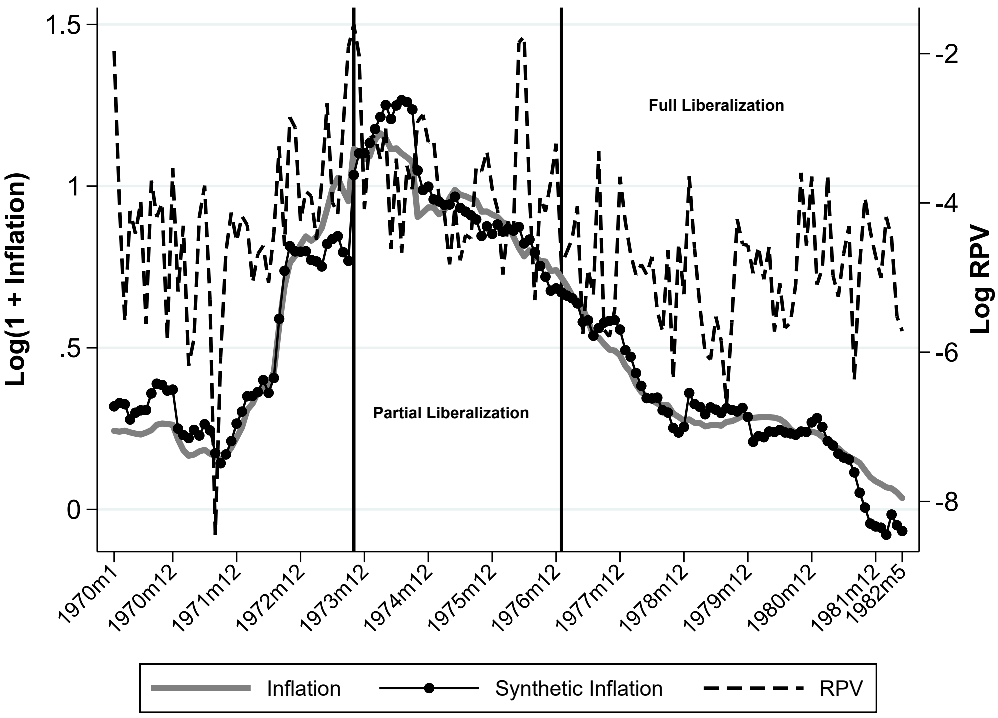

Table 1. Relative Price Variability–Inflation Relationship, Classical Measure of RPV

VARIABLES	(1)	(2)	(3)
Annual Inflation (π_t^a)	0.020***	0.024**	0.026**
	(0.01)	(0.01)	(0.01)
$D_t^{73} \times \pi_t^a$		-0.006	-0.008
		(0.01)	(0.01)
$D_t^{77} \times \pi_t^a$		-0.021**	-0.021**
		(0.01)	(0.01)
Lagged RPV	0.385*	0.381*	0.394*
	(0.21)	(0.20)	(0.20)
Observations	149	149	149
Macroeconomic Control Variables			✓

Note: Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%. Results estimated from a regression of the form: $RPV_t = \gamma_0 + \alpha_1\pi_t + \alpha_2 D_t^A \pi_t + \alpha_3 D_t^B \pi_t + \mathbf{Z_t}\boldsymbol{\zeta} + \varepsilon_t$, where RPV_t is the relative price variability at time t, measured as in Fischer et al. (1981) and Parks (1978), π_t is the annual inflation rate, $\mathbf{Z}_t$ correspond to the macroeconomic variables plus a lag on the dependent variable, D_t^A is equal to one between October 1973 and December 1976, and zero otherwise and D_t^B is equal to one between January 1977 and May 1982, and zero otherwise.

Table 2. Relative Price Variability–Inflation Relationship, Caraballo and Dabús (2013) RPV measure

VARIABLES	(1)	(2)	(3)
Annual Inflation (π_t^a)	0.011***	0.015*	0.017*
	(0.00)	(0.01)	(0.01)
$D_t^{73} \times \pi_t^a$		-0.006	-0.008
		(0.01)	(0.01)
$D_t^{77} \times \pi_t^a$		-0.016**	-0.016**
		(0.01)	(0.01)
Lagged RPV	0.440**	0.432**	0.453**
	(0.18)	(0.18)	(0.18)
Observations	149	149	149
Macroeconomic Control Variables			✓

Note: Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%. Results estimated from a regression of the form: $RPV_t = \gamma_0 + \alpha_1\pi_t + \alpha_2 D_t^A \pi_t + \alpha_3 D_t^B \pi_t + \mathbf{Z_t}\boldsymbol{\zeta} + \varepsilon_t$, where RPV_t is the relative price variability at time t, measured as in Caraballo and Dabús (2013), π_t is the annual inflation rate, $\mathbf{Z}_t$ correspond to the macroeconomic variables plus a lag on the dependent variable, D_t^A is equal to one between October 1973 and December 1976, and zero otherwise and D_t^B is equal to one between January 1977 and May 1982, and zero otherwise.

Table 3. Relative Price Variability–Inflation Relationship, Using log adjustment in inflation

VARIABLES	(1) Classic	(2) Classic	(3) CD	(4) CD
$\log(1+\pi_t^a)$	0.040**	0.042**	0.026*	0.028*
	(0.02)	(0.02)	(0.01)	(0.01)
$D_t^{73} \times \log(1+\pi_t^a)$	-0.008	-0.010	-0.008	-0.011
	(0.02)	(0.02)	(0.01)	(0.01)
$D_t^{77} \times \log(1+\pi_t^a)$	-0.031**	-0.031**	-0.023**	-0.023**
	(0.01)	(0.01)	(0.01)	(0.01)
Lagged RPV	0.385*	0.398*		
	(0.21)	(0.21)		
Lagged RPV			0.432**	0.452**
			(0.18)	(0.18)
Observations	149	149	149	149
Macroeconomic Control Variables		✓		✓

Note: Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%. Results estimated from a regression of the form: $RPV_t = \gamma_0 + \alpha_1\pi_t + \alpha_2 D_t^A log(1+\pi_t) + \alpha_3 D_t^B log(1+\pi_t) + \mathbf{Z_t}\boldsymbol{\zeta} + \varepsilon_t$, where π_t is the annual inflation rate, $\mathbf{Z}_t$ correspond to the macroeconomic variables plus a lag on the dependent variable, D_t^A is equal to one between October 1973 and December 1976, and zero otherwise and D_t^B is equal to one between January 1977 and May 1982, and zero otherwise.

Table 4. Results Using Synthetic Total-RPV

VARIABLES	(1)	(2)	(3)	(4)	(5)
Annual Inflation (π_t^a)	0.013**	0.014***	0.014***		
	(0.01)	(0.00)	(0.00)		
$D_t^{73} \times \pi_t^a$		-0.003	-0.002		
		(0.00)	(0.00)		
$D_t^{77} \times \pi_t^a$		-0.008***	-0.008***		
		(0.00)	(0.00)		
$\log(1+\pi_t^a)$				0.021***	0.020***
				(0.01)	(0.01)
$D_t^{73} \times \log(1+\pi_t^a)$				-0.005	-0.004
				(0.01)	(0.01)
$D_t^{77} \times \log(1+\pi_t^a)$				-0.012***	-0.012***
				(0.00)	(0.00)
Lagged RPV	0.844***	0.849***	0.854***	0.880***	0.887***
	(0.06)	(0.07)	(0.07)	(0.06)	(0.05)
Observations	149	149	149	149	149
Macroeconomic Control Variables			✓		✓

Note: Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%. Results estimated from a regression of the form: $RPV_t = \gamma_0 + \alpha_1\pi_t + \alpha_2 D_t^A\pi_t + \alpha_3 D_t^B\pi_t + \mathbf{Z_t}\boldsymbol{\zeta} + \varepsilon_t$, where π_t is the annual inflation rate, $\mathbf{Z}_t$ correspond to the macroeconomic variables plus a lag on the dependent variable, D_t^A is equal to one between October 1973 and December 1976, and zero otherwise and D_t^B is equal to one between January 1977 and May 1982, and zero otherwise.

Table 5. Results Using Within Relative Price Variability (WRPV), 1970-1976

VARIABLES	(1)	(2)	(3)	(4) AB Estimates	(5) AB Estimates
Annual Inflation $[\phi_1]$	0.074***	0.052***	1.878**	0.013***	0.055***
	(0.01)	(0.01)	(0.78)	(0.00)	(0.01)
$D_{it}^{73} \times \pi_t^a$ $[\phi_4]$		-0.212**	-0.212**	-0.019***	-0.011*
		(0.08)	(0.08)	(0.01)	(0.01)
Observations	1,932	1,932	1,932	1,886	1,886
Product FE	✓	✓	✓	✓	✓
Month FE			✓		✓
Year FE			✓		✓

Note: : Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%. D_t is equal to one after establishing the partial liberalization, and zero otherwise, D_i is equal to one if the product was liberalized, and zero otherwise and $D_{it} = D_i \times D_t$.

Table 6. Results Using Within Relative Price Variability (WRPV), 1973-1982

VARIABLES	(1)	(2)	(3)	(4) AB estimates	(5) AB estimates
Annual Inflation $[\phi_1]$	0.092***	0.062	-0.111	0.011***	0.038***
	(0.01)	(0.06)	(0.13)	(0.00)	(0.00)
$D_{it}^{77} \times \pi_t^a$ $[\phi_4]$		-0.256*	-0.256*	-0.026***	-0.020***
		(0.12)	(0.13)	(0.01)	(0.01)
Observations	2,369	2,369	2,369	2,369	2,369
Product FE	✓	✓	✓	✓	✓
Month FE			✓		✓
Year FE			✓		✓

Note: : Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%. D_t is equal to one after established the total liberalization, and zero otherwise, D_i is equal to one if the product was liberalized, and zero otherwise and $D_{it} = D_i \times D_t$.

Table 7. Nested Model, by Product, 1970-1982

VARIABLES	(1)	(2)	(3) AB Estimates
Annual Inflation	0.078***	0.052***	0.014***
	(0.01)	(0.01)	(0.00)
$D_{it}^{73} \times \pi_t^a$		-0.212**	-0.019***
		(0.08)	(0.01)
$D_{it}^{77} \times \pi_t^a$		-0.256*	-0.022***
		(0.12)	(0.01)
Observations	3,427	3,427	3,381
Product FE	✓	✓	✓

Note: : Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%.

Table 8. Nested Model, by Product, 1970-1982 (Including Time Effects)

VARIABLES	(1) 70 – 76	(2) 73 – 82	(3) 70 – 82	(4) 70 – 76 (AB)	(5) 73 – 82 (AB)	(6) 70 – 82 (AB)
$D_{it}^{73} \times \pi_t^a$	-0.212**		-0.212**	-0.011*		-0.014***
	(0.08)		(0.08)	(0.01)		(0.01)
$D_{it}^{77} \times \pi_t^a$		-0.245*	-0.256*		-0.020***	-0.016**
		(0.12)	(0.13)		(0.01)	(0.01)
Observations	1,932	2,599	3,427	1,886	2,369	3,381
Product FE	✓	✓	✓	✓	✓	✓
Month × Year FE	✓	✓	✓	✓	✓	✓

Note: : Robust standard errors are presented in parentheses. Significance levels: * 10%, ** 5%, *** 1%.