Regional inflation dynamics using space-time models

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Abstract

This paper provides empirical evidence of the role of spatial factors on the determination of inflation dynamics for a representative set of tradable commodities in Chile. We present a simple model that explains inflation divergence across regions in a monetary union with similar preferences as a consequence of the geographical allocation of producers in the different regions. Our results indicate that spatial allocation together with transport costs are important determinants of regional inflation while macroeconomic common factors do not play an important role in this process. Existing literature had obtained the opposite result for Europe and the reasons for that difference warrant further investigation. Moreover, we find that geographical distance seems to be a more appropriate measure of neighbourhood than the adjacent of regions.

Keywords: *regional inflation dynamics, space-time models, Chile*

JEL codes: *E31, E52, E58, R11, C23, C21*

1. Introduction

Explaining persistent inflation differentials across the various geographical areas that make up a monetary union has been a recurrent topic in the economic literature. Two recent and important contributions in this field are Altissimo et al (2005), who present a theoretical model to explain inflation dispersion in the non traded sector, and Andres et al. (2008) who focus on tradable goods, suggesting that inflation differentials may be substantial over the business cycle mainly because of different preferences of individuals in different countries.

In this paper we show that another factor - transport costs - can explain persistent inflation differentials for tradable goods even when individual preferences in the different regions are identical. This intuition is motivated with a simple modified version of the Obstfeld and Rogoff (1995) model (O&R henceforth) for two different regions in a monetary union. Unlike O&R, in our model exchange rates are fixed and do not adjust to fulfill the law of one price. Instead, under sticky prices and in the presence of transport costs, the proportion of producers in the two regions can explain the asymmetric reactions of regional prices and incomes to the same type of macro shocks and thus explain persistent inflation divergence across the monetary union.

The role of spatial factors in the determination of inflation dynamics is tested for a representative set of 98 tradable commodities, whose prices have been taken monthly for 23 cities of Chile in the period 2003:01-2006:09. Due to its natural geography and climate that prevent perfect price arbitrage, the Chilean case allows a natural application of spatial econometric models to the explanation of the heterogeneity of inflation dynamics at the regional and product level.

This paper introduces two important novel features with respect to previous work that tested spatial price homogeneity (law of one price), such as Parsley and Wei (1996) and Ceccheti et al. (2002) for the U.S., or Beck et al. (2009) for the Euro Area. First, as far as we are aware, ours is the first attempt to investigate the heterogeneity of inflation dynamics for an emerging market with such a level of detail. Indeed, we explore the product and geographical dimensions of Chilean inflation using spatial econometric models. The second aspect relates to our study of inflation dispersion for individual prices and not for price indices. This is an important issue given that a price index could evolve differently across regions just because of different weights in the representative basket of consumption. The individual consideration of homogeneous product categories eliminates this problem.

Our results indicate an important degree of spatial correlation in the determination of commodity prices, supporting the theoretical result that persistent inflation differences across space can be due to the geographical allocation of producers. Also, in contrast to what Beck et al. (2009) have found for the Euro Area, in Chile common macroeconomic factors only explain a small proportion of the variability of inflation for the different commodities.

The structure of this paper is as follows. Section 2 proposes a simple model that explains regional divergence in the inflation rates through the role of transport costs. The following section presents the dataset used in the empirical analysis and explains some of its features. Section 4 discusses the most important empirical results obtained from the estimation of a range of spatial econometric models. Some concluding remarks follow in Section 5.

2. Theoretical underpinnings

The role of transport costs in determining inflation dynamics can be explained theoretically by a simplified version of the O&R model for different regions in a single country. An important feature of this framework is that all goods are traded but imported goods are subject to a transport cost. Unlike O&R who focus their attention on the effect of monetary policy and the exchange rate adjustment to maintain the purchasing power parity between prices in two different countries, here we deal with regions inside the same country and the relationships between prices in each region are governed by transport costs instead of exchange rates. Under the assumption of sticky prices, shocks to demand and transport costs can alter relative prices and generate asymmetric reactions across regions, even if all individuals have the same preferences independently of their location.

We assume the world is inhabited by a continuum of individual monopolistic producers, indexed by $z \in [0,1]$, each of whom produces a single differentiated good, also indexed by z. All producers locate in one of two regions, central or peripheral. Central regions produce in the interval $[0, n]$, whereas peripheral regions are located in $(n, 1]$. Each agent located in one of the regions produces a variety of one type of good, $y(z)$, in which that region is specialized. Independently of their location, all producers sell some of their production in the central market, $y_{C,t}(z)$, and the rest in the peripheral market, $y_{P,t}(z) = y(z) - y_{C,t}(z)$. Because each variety is unique, they enjoy some monopolistic power at both their home region and outside.

All individuals throughout the country have identical preferences over a consumption index, real money balances and effort expended in production, whether they locate in the central or in the peripheral region. The intertemporal utility function of a typical agent j is given by

$$U_t^j = \sum_{s=t}^{\infty} \beta^{s-t} \left[\log C_s^j + \chi \log \frac{M_s^j}{P_s} - \frac{\kappa}{2} y_s(j)^2 \right] \quad (2.1)$$

The variable C is a real consumption index

$$C^j = \left[\int_0^1 c^j(z)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta}{\theta-1}} \quad (2.2)$$

where $c^j(z)$ is the jth Home individual's consumption of good z, and $\theta > 1$.

Let $p(z)$ be the Home-currency price of good z. Then the Home money price index is

$$P = \left[\int_0^1 p(z)^{1-z} dz \right]^{\frac{1}{1-\theta}} \quad (2.3)$$

The most important difference between this framework and the O&R model is the derivation of the relationship between P and P^*. Here, it is assumed that all goods can be traded, although in doing so transport costs are incurred according to an iceberg transport technology (see Fujita et al. 1999). Then, if the transport cost of a good from one region to another is T_t, the relationship between the prices in the two regions is given by

$$p(z)T_t = p^*(z) \quad if\ z \leq n$$
$$p(z) = T_t p^*(z) \quad if\ z > n \quad (2.4)$$

where $T_t > 1$.

Now, we can write the central price index as

$$P = \left[\int_0^1 p(z)^{1-z} dz\right]^{\frac{1}{1-\theta}} = \left[\int_0^n p(z)^{1-\theta} dz + \int_n^1 [T_t p^*(z)]^{1-\theta} dz\right]^{\frac{1}{1-\theta}} \quad (2.5)$$

Similarly, the peripheral price index is

$$P^* = \left[\int_0^1 p^*(z)^{1-z} dz\right]^{\frac{1}{1-\theta}} = \left[\int_0^n [p(z)T_t]^{1-\theta} dz + \int_n^1 p^*(z)^{1-\theta} dz\right]^{\frac{1}{1-\theta}} \quad (2.6)$$

An important point to note is that, although all agents have similar preferences, the law of one price is not necessarily met. In general:

$$P \neq P^* \quad (2.7)$$

Prices in the central and peripheral regions will be different if the location of agents in the two regions is asymmetric. In this model, the proportion of individuals allocated to the different regions is considered exogenous as the location decision could be affected by geographical, political, economic and historical reasons. Hence we can suppose that in general the allocation of producers will in fact be asymmetric. As we show later, in this case transport costs play a key role in perpetuating price dispersion.

Similarly to O&R, the only internationally traded asset is a riskless real bond denominated in the composite consumption good. The period budget constraint for a representative Home individual j can be written in nominal terms as

$$P_t B_{t+1}^j + M_t = P_t(1+r_t)B_t^j + M_{t-1} + p_t(j)y_t(j) - P_t C_t^j \quad (2.8)$$

where r_t denotes the real interest rate on bonds between $t-1$ and t, $y_t(j)$ is output for good j and $p_t(j)$ is the domestic currency price. The variable M_{t-1}^j is agent j's holdings of nominal money balances entering at period t.

The economy is closed and we do not consider a foreign country as our interest lies in analyzing the differences across regions in the same country. Therefore, a single monetary aggregate is considered for both regions. Under the assumption that the monetary authority runs a balanced budget at each period and given that, for simplicity, we do not consider taxes and government spending, the following condition must be verified

$$0 = \frac{M_t - M_{t-1}}{P_t} \quad (2.9)$$

The set of equilibrium conditions obtained by following a similar approach to O&R is confined to Appendix 1. Given that prices are determined by monetary policy and both regions face the same type of monetary policy shocks the only value of $\hat{t}_t$ that is consistent with similar prices in the central and peripheral areas is $\hat{t}_t = 0$. However, it is far more interesting to study the case in which transport costs are altered and the adjustment of prices to the new level of transport costs takes place with a one-period lag. Movements in transport costs could be explained either by a common demand shock that affects demand for all goods in the economy or by an international oil shock. In both cases, shocks at the national level will exert an asymmetric effect in the two regions. To see this, notice that, under sticky producer prices, any change in transport costs would alter consumer prices in the central and peripheral areas in the following way:

$$\hat{p} = (1-n)\hat{t} \quad (2.10)$$

$$\hat{p}^* = n\hat{t} \quad (2.11)$$

If $n > 0.5$ the consumer price index after the shock will be higher in the peripheral region compared to the central region. This happens because individuals in the peripheral region have to pay for the cost of transporting all goods produced in the most populated areas (the central region). The asymmetry in consumer price indices also has an asymmetric effect on the level of consumption and income in the two areas, as given by the following equations:

$$\hat{y} - \hat{y}^* = \theta[-n\hat{t} + (1-n)\hat{t}] \quad (2.12)$$

$$\hat{y} - \hat{y}^* = -\frac{\theta}{1+\theta}(c - c^*) \quad (2.13)$$

Therefore, a shock to transport costs alters the distribution of income in the two regions. Moreover, by subtracting the Euler equations (2.12) and (2.13) in the central and peripheral regions it can be seen that these relative changes in consumption levels are always permanent.

To sum up, it is clear from this model that the existence of transport costs effectively prevents the elimination of regional inflation differentials as changes in transport costs will lead to permanent changes in relative consumer prices in the two regions. Accordingly, we should find in the empirical analysis that regional inflation is not only determined by monetary policy but also by transport costs, which in turn are a function of the distance across production locations. Testing this theory for the case of Chile is the main task of the subsequent sections.

3. Product inflation data

For our analysis we collected a panel of prices covering an important range of different types of foods and drinks as well as oil products, summing up to a total of 98 different products. Price data was taken on a monthly basis in the period 2003:01-2006:09 for 23 cities that are representative of the 11 regions in Chile. Chile has an unusual ribbon-like shape which is on average 175 kilometers wide and 4,300 kilometers long. This length is higher than, for example, the distance from Madrid to Moscow (3,438 Km) - see Figure 1. A detailed description of the different sectors and regions considered is confined to Appendix 2.

The data are freely available from the National Statistical Institute of Chile ("Instituto Nacional de Estadística") at the URL http://www.ine.cl. This institution stopped publishing information on regional prices after September 2006 and so more recent data cannot be collected. However, even if this information had been available, inflation dynamics after that date followed a pattern that was not consistent with its equilibrium values in equilibrium and would represent an important break in the panel. More specifically, due to the higher increase in world food and oil since 2006, the average annual rate of Chilean inflation (i.e. the increase of the general index of consumer prices) was 7.8% and 7.1% in 2007 and 2008 respectively, while it had oscillated between 2% and 3% in the period 2004-2006.

Inflation rates for each of the items (π_t) are computed as year-on-year percentage changes in the price index in the following way:

$$\pi_t = 100 * \left(\frac{P_t - P_{t-12}}{P_{t-12}}\right) \quad (3.1)$$

where P_t denotes the respective product price in a given region.

Compared to other related papers such as Cecchetti et al. (1999), Beck et al. (2009) and Tena et al. (2009), an important advantage of our database is that we are considering individual prices instead of disaggregate price indices that include a basket of products even at the disaggregate level. These indices could evolve differently simply because of different regional tastes for the items in the consumption basket and not because of the different dynamics of prices in the different regions.

There are no observations for item 28 (fish) in the city of Punta Arenas and therefore we exclude information from this city for that product. Besides, the panel contains a small number of missing values that represent about 0.5% of the total number of observations. We tackle such data irregularities in a factor model framework by using the EM algorithm together with PC decomposition (see for example Stock and Watson (2002) and Schumacher and Breitung (2008)). More specifically, using the inflation information available for the 23 cities, we estimate the most important common factors for governing inflation in each of the 98 items (except product 28 whose observations are available for only 22 cities). Then, in a second step, the regression of each of the individual inflation series on the common factor is used to complete the missing values. The EM algorithm repeats steps 1 and 2 until convergence.

For a formal test on the number of unit roots in the panel we follow Parsley and Wei (1996) by using the panel unit root test proposed by Levin et al. (2002). More precisely, for each of the 98 items, the basic regression specification is

$$\Delta\pi_{k,t} = \mathrm{c_k} + \beta\pi_{k,t-1} + \sum_{i=1}^{p} \gamma_i \, \Delta\pi_{k,t-1} + \varepsilon_t \qquad (3.2)$$

where $\pi_{k,t}$ is the annual growth rate of prices in city k at time t; c_k is the constant term specific to the kth city (i.e. we have a series of 23 dummies); and ε_t is the error term.

According to Levin et al. (2002), the critical values for T=25 and N=25 (that is, approximately our panel size) at the 1 and 5% significance levels are -8.27 and -7.74, respectively. The results of this test indicate that the null hypothesis of non stationarity could be rejected for 70% of the commodities at the 5% level. Thus, inflation can be considered as being generated by a stationary process in most cases but not all. Indeed, there is an open debate in the literature on whether inflation is stationary or generated by a unit root process (see for example Culver and Papell (1997) for a discussion on this issue). Assuming that there is not an equilibrium rate for inflation is quite a strong hypothesis. But, it is also true that the mean level of inflation is typically affected by different stochastic breaks and to approximate these sporadic breaks with a unit root can be, in some circumstances, a good approach.

According to this analysis we consider inflation as a stationary process and test the impact of spatial variables on its evolution. We also check the robustness of our results even if inflation is not stationary.

4. Econometric specifications and results

We initially estimate the following equation for each of the 98 commodities in the sample:

$$\pi_{k,t} = \alpha_k + \beta_1 \pi_{k,t-1} + \beta_2 f_t + \beta_3 f_{t-1} + \varepsilon_{k,t} \quad (4.1)$$

where $\pi_{k,t}$ is the annual inflation rate for the *k-th* city at time t; f_t is a national common factor; and $\varepsilon_{k,t}$ is the error term. This specification is denoted as model M1.[2]

Note that equation (4.1) resembles the one proposed by Beck *et al.* (2009) in the sense that it considers the influence of national common factors at the national level on the dynamics of regional inflation. Indeed, to make our results comparable to Beck *et al.* (2009), we estimate common factors based on national macroeconomic variables such as the Chilean short-term interest rate, unemployment, the growth rate of oil prices, Chilean money supply, the nominal effective exchange rate, unit labour costs and industrial production.[3] However, unlike them, we only consider a common factor at the national level and not at the continental level because there is not a common monetary policy for all South American countries. A summary[4] of the estimation results - exhibited in the first column of Table 1 for the significance of each variable and in the first column of Table 2 for the explanatory power of each variable - shows that, for most commodities the dynamics of inflation is determined by its own past values and not by the common factors.

We then test for the presence of spatial correlation in the residuals of the model by defining a weights matrix that takes positive values for cities in the same region and adjacent regions and zero otherwise. We do this by defining a spatial lag order as:

$$L^{(1)}\pi_i = \sum_{\substack{j \in G_s \\ j \neq i}} w_{ij}^{(1)} \pi_j \quad (4.2)$$

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[2] The model includes the lagged dependent variable, which is potentially endogenous. However, using Monte Carlo simulations, Beck and Katz (2004) find that the nickel bias is low (2% or less) once $T = 20$, and they advise the use of a least-square estimator with a lagged dependent variable included if T is at least 20. Our sample contains 45 months, hence we do not correct for endogeneity of the lagged dependent variable.

[3] Appendix 2 contains a detailed description of these variables.

[4] Due to the large number of commodities used, only a summary of results is presented. The full set of results is available from the authors.

where G_s is the set of s neighbours of order (1).

In this case, the weights $w_{ij}^{(1)}$ depend on the number of cities in the different regions. For example, if for a certain location, there are 5 different cities in the same region and adjacent regions, $w_{ij}^{(1)} = 1/5$ for each of the 5 cities and 0 for the remaining ones. Therefore, the following properties are met: 1) $w_{ij}^{(s)} \geq 0$; 2) $w_{ii}^{(s)} = 0$; and 3) $\sum_{\substack{j \in G_s \\ j \neq i}} w_{ij}^{(s)} = 1$ (see, for example, Anselin (1988) and Arbia (2006)).

We carried out several tests of the presence of spatial correlation in the residuals of model (4.1), such as Moran's I-statistic, the likelihood ratio test, the Wald test and the Lagrange multiplier test. The results of all the aforementioned tests indicate that the null of no spatial correlation could be rejected in more than 60% of the commodities at the 5% significance level.[5] These results suggest that specification (4.1) could be improved by taking into account the interrelations of each city with other cities in the same and adjacent regions. Therefore, we augment the previous model by considering the following specification

$$\pi_{k,t} = \alpha_k + \beta_1 \pi_{k,t-1} + \beta_2 L^{(1)} \pi_{k,t} + \beta_3 L^{(1)} \pi_{k,t-1} + \beta_3 f_t + \beta_4 f_{t-1} + \beta_5 \pi_t + \beta_6 \pi_{t-1} + \varepsilon_{k,t} \quad (4.3)$$

where π_t is the annual growth rate of the Chilean consumer price index. The augmented model is denoted by M2.

This specification allows us to take into account spatial correlation in the inflation rates of the different cities and also the influence of the general inflation rate for each product. Note however that the likelihood function for model (4.3) cannot be maximized analytically due to the high degree of nonlinearity in the parameters. Therefore, we approximate this estimation by making use of the pseudo-likelihood procedure in LeSage (1999).

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[5] See Anselin (1988) for a definition of the weights matrix in spatial econometric models and for different tests of spatial correlation.

Space-dependency is found in 81% of the commodities, whilst the common factor determines inflation for only 10% of them at the 5% level (Table 1, second column). On average, the common factor explains 16% of the observed inflation variance, with the inflation of neighbour regions explaining 7% contemporaneously and 21% with a one period lag (Table 2, second column). This result confirms the theoretical hypothesis of section 2 according to which distance (transport cost) determine prices contemporaneously to some extent, but with a one-period lag to a higher extent. On the other hand, our empirical results are at odds with what Beck et al (2009) had found for Europe, where the macroeconomic common factor was the most important determinant of inflation dynamics. Moreover, the finding that neighbours matter more than common factors (even if with a time lag) justifies the introduction of spatially lagged variables in models of determination of regional inflation.

The fact that a macroeconomic common factor is not an important variable to describe inflation dynamics in many commodities does not necessarily mean that an important part of the dispersion observed in the various commodity prices is not affected by a common shock at the national level. It could be that case that there are shocks specific to each commodity. This possibility is tested by obtaining a national common factor for each of the commodities from the inflation dynamics for that item in the 23 cities and estimating a model similar to (4.3), which we denote as M3. The consideration of commodity-specific common factors improves our econometric specifications. In fact, commodity-specific common factors are significant in 91% of the regressions compared to the 10% found in the previous specification (see third column of Table 1) and explain 31% of inflation dynamics compared to 16% previously (see third column of Table 2). However, also in this case spatial variables have a significant impact on the determination of prices for 62% of the commodities and do not lose explanatory power.

One potential drawback of our empirical results is the ad hoc consideration of the weights matrix. In fact, the choice of weights is typically a discretionary decision made by the researcher and a different selection of neighbours could result in different conclusions. To deal with this problem we test whether the residuals of the previous model still contain some degree of spatial correlation that could be captured by a different contiguity matrix. In particular, we consider a second order of contiguity by defining regions that are neighbours of neighbours, $L^{(2)}$, and a weights matrix based on kilometre distances instead of adjacent regions. That is, in the latter case w_{ij} is defined as the distance from city i to j divided by the sum of distances from i to all the other 22 cities. We find that the second order of contiguity is significant in only 17% of the residuals while distance is significant in 60% of them at the 5% significance level.

Given this result, a model similar to model M3 (model M4) was estimated using a contiguity matrix based on distance. Spatial variables became significant in 76% of the commodities under this specification (fourth column of Table 1) and maintain the same explanatory power (fourth column of Table 2), which suggests that, consistently with our theory, it is geographical distance and not administrative boundaries that influences inflation in the different cities.

As a robustness exercise, we estimate a model similar to M4 by GMM (denoted as M5).[6] Under GMM estimation, spatial variables exert an even more important role on the determination of inflation for the different commodities. Contemporaneous and lagged inflation in nearby cities are now significant for respectively 53% and 44% of the commodities (fifth column of Table 1) and they jointly explain 48% of inflation variance (fifth column of Table 2).

Our final robustness exercise considers a similar estimation to M4 using variables in first differences (model M6). This specification makes sense given that, although the unit root tests performed in the previous section suggest that inflation is a stationary process, the evidence is not compelling (it was found for around 70% of the cases). The results of the estimation of model M6 confirm that spatial variables also play an important role on the determination of the first differences of inflation for the various commodities (sixth columns of Tables 1 and 2).

It is possible to conclude from this analysis that consistently with our theory spatial factors play a key role in the determination of Chilean inflation. Moreover, geographical distance seems to be a more appropriate measure of neighbourhood than the adjacency of regions. Common macroeconomic factors do not explain an important proportion of commodity price dynamics; however national factors that are specific to each commodity turn out to be important explanatory variables.

[6] The instruments used in the GMM estimation are: common factor (contemporaneous and first lag); national inflation (contemporaneous and two first lags); neighbour inflation (up to the fifth lag); dependent variable (second to fifth lags). The results are robust to the selected lags.

5. Concluding remarks

We have analysed the determinants of inflation for 98 commodities in 23 major Chilean cities. The results obtained indicate that inflation differentials can be observed across regions that are affected by the same type of macroeconomic shocks. One possible explanation for this finding is the key role played by geographical distance in the determination of inflation rates for the different cities. These empirical results are consistent with a model with sticky prices where the location of firms in the different regions of a country together with transport costs could explain inflation divergence that is persistent through time.

Future lines of research are suggested by this work. First, it would be interesting to augment the theoretical model to explain endogenously the evolution of transport costs. In this way, it should be possible to analyse the implications of different types of worldwide economic shocks at the regional level. The second extension relates to the analysis of monetary policy. Given that inflation could diverge across regional areas, central bankers should consider the various sources of inflation heterogeneity in order to conduct an optimal monetary policy. This is especially relevant for the Chilean case where the overall inflation rate is computed based only on prices in the capital (Santiago de Chile). Finally, it is important to understand why the relative importance of common macroeconomic shocks and of transport costs in the determination of regional inflation dynamics varies in different parts of the world. Starting from the cases of Chile studied in this paper and of the Euro Area (Beck et al 2009), the investigation of this issue for other countries is warranted in order to determine under which circumstances regional inflation dynamics is more dependent on macroeconomic shocks or on transport costs. This is an important issue because the first factor can be influenced by monetary policy, whilst the latter cannot.

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__Appendix 1__

__First order conditions in the theoretical model__

The equilibrium (in log-linear form) is represented by the following system of equations:

$$\hat{p}_t = n\hat{p}_t(c) + (1-n)[\hat{t}_t + \hat{p}_t^*(p)] (A1)$$

$$\hat{p}_t^* = n[\hat{p}_t(c) + \hat{t}_t] + (1-n)\hat{p}_t^*(p) (A2)$$

$$\hat{p}_t - \hat{p}_t^* = -n\hat{t}_1 + (1-n)\hat{t}_t (A3)$$

$$\hat{y}_t = \theta[\hat{p}_t - \hat{p}_t(c)] + \hat{c}_t^w \ (A4)$$

$$\hat{y}_t^* = \theta[\hat{p}_t^* - \hat{p}_t^*(p)] + \hat{c}_t^w \ (A5)$$

$$\hat{c}_{t+1} = \hat{c}_t + \frac{\delta}{1+\delta}\hat{r}_{t+1} \ (A6)$$

$$\hat{c}_{t+1}^* = \hat{c}_t^* + \frac{\delta}{1+\delta}\hat{r}_{t+1} \ (A7)$$

$$\widehat{m}_t - \hat{p}_t = \hat{c}_t - \frac{r_{t+1}}{1+\delta} - \frac{\hat{p}_{t+1} - \hat{p}_t}{\delta} \ (A8)$$

$$\widehat{m}_t - \hat{p}_t^* = \hat{c}_t^* - \frac{r_{t+1}}{1+\delta} - \frac{\hat{p}_{t+1}^* - \hat{p}_t^*}{\delta} \ (A9)$$

$$\bar{c} = \delta\bar{b} + \bar{p}(c) + \bar{y} - \bar{p} \ (A10),$$

$$\bar{c}^* = -\left(\frac{n}{1-n}\right)\delta\bar{b} + \bar{p}^*(c) + \bar{y}^* - \bar{p}^* \ (A11)$$

where for each variable x_t, we define $\hat{x}_t \equiv dx_t/x_0$ and $\bar{x}$ corresponds to its value in equilibrium.

Equations (A1) and (A2) are the log-linear form of the central and peripheral price index under the assumption of asymmetry among each region's producer and the relationship between prices in the two regions are sketched in equation (A3). The log-linear form for the demands of an individual good produced in the central and peripheral regions are described in equations (2.14) and (2.15) in which we define world consumption as

$$\hat{c}_t^w = n\hat{c}_t + (1+n)\hat{c}_t^* = n\hat{y}_t + (1+n)\hat{y}_t^* = \hat{y}_t^w \quad (A12)$$

Equations (A4) to (A9) express the first order conditions from the maximization of the individual utility function whereas the last two equations comes from the integration of the individual's period budget constraint over time and the imposition of the transversality condition.

Appendix 2

Time Series

The time series considered in the analyses can be freely obtained from the Chilean National Statistical Institute at the following URL: http://www.ine.cl. *The panel data base consists of observations for 98 different items in 23 different Chilean cities on monthly basis for the period 2003:01-2006:09.*

The cities and items in the sample are depicted below

Cities

Chillán	*Coihaique*	*Concepción*	*Curicó*
Linares	*Los Ángeles*	*Osorno*	*Puerto Montt*
Rancagua	*San Antonio*	*San Fernando*	*Talca*
Temuco	*Valdivia*	*Valparaiso*	*Antofagasta*
Arica	*Copiapo*	*Iquique*	*La Serena*
Quillota	*Punta Arenas*	*Los Andes*	

Items

r1: Normal bread (kg)	***r2***:Special bread (no package) (kg)	***r3***:Rice (kg)
r4:Flour (kg)	***r5:*** Oats (500 g)	***r6:*** Noodles Nº 5 (400 g)
r7: Noodles Nº 87 (400 g)	***r8:*** Spiral Noodles (400 g)	***r9:*** Quifaro Noodles (400 g)
r10: Wafer biscuit (140 g)	***r11:*** Lemon biscuit (140 g)	***r12:*** Water biscuit (210 g)
r13: Salted potatoes (230 g)	***r14:*** Olives (300 g)	***r15:*** Pai (15 persons)
r16: Meat (best quality) (kg)	***r17:*** Beef ribs (kg)	***r18:*** Rump, Cap and Tail Off (kg)
r19: Filet (kg)	***r20:*** Sirloin Tip (kg)	***r21:*** Shank (kg)
r22: Minced meat 10 % fat (kg)	***r23:*** Pork chop (kg)	***r24:*** Pork rib cage (no seasoning) (kg)
r25: Chicken (kg)	***r26:*** Chicken breast (kg)	***r27:*** Turkey breast (kg)
r28: Fish (kg)	***r29:*** Canned mackerel	***r30:*** Canned tuna (184 g)

	(425 g)	
r31: Canned sardines (125 g)	***r32:*** Ham (kg)	***r33:*** *Culin* bologna (kg)
r34: Sausages (20 units)	***r35:*** Spicy sausages (kg)	***r36:*** Beef Paté (125 g)
r37: Mayonnaise (250 cc)	***r38:*** Eggs (12 units)	***r39:*** Milk (bag) (lt)
r40: Milk (pack) (lt)	***r41:*** Powdered milk (1,6 kg)	***r42:*** Powdered milk (kg)
r43: Sweetened condensed milk (400 g)	***r44:*** Cereal (box) (510 g)	***r45:*** Salted butter (kg)
r46: Cheese (kg)	***r47:*** Cream Cheese (kg)	***r48:*** Cheese (bag) (360 g)
r49: Yogurt (175 g)	***r50:*** Powered gelatine (160 g)	***r51:*** Powered caramel pudding (180 g)
r52: Vegetable oil (lt)	***r53:*** Sunflower oil (lt)	***r54:*** Salted Margarine (250 g)
r55: Avocado (kg)	***r56:*** Organic tomato (kg)	***r57:*** Normal tomato (Kg)
r58: Lemons (kg)	***r59:*** Apples (kg)	***r60:*** Oranges (Kg)
r61: Bananas (kg)	***r62:*** Canned peaches (590 g)	***r63:*** Canned peas (310 g)
r64: Potatoes (kg)	***r65:*** Garlics (3 units)	***r66:*** Onions (kg)
r67: Lettuce (one)	***r68:*** White cabbage (one)	***r69:*** Carrots (bunch)
r70: Pumpkin (kg)	***r71:*** Lentils 5 mm (kg)	***r72:*** Beans (kg)
r73: Green beans (kg)	***r74:*** Tomato sauce (bottle) (250 g)	***r75:*** Tomato sauce (tetra) (215 g)
r76: Sugar (kg)	***r77:*** Marmalade (250 g)	***r78:*** Salt (kg)
r79: Instantaneous soup (70 g)	***r80:*** Chicken gravy cubes (8 units)	***r81:*** Coffee (170 g)
r82: Fortifier for milk (400 g)	***r83:*** Tea (250 g)	***r84:*** Tea bags (20 units)
r85: Bottled soft drink (2 lt)	***r86:*** Canned soft drink (355 cc)	***r87:*** Organic juice (lt)
r88: Powder juice (45 g)	***r89:*** Ice cream (lt)	***r90:*** Wine (lt)
r91: Sparkling mineral water (1,6 lt)	***r92:*** Bottled beer (lt)	***r93:*** Canned beer (355 cc)
r94: Pisco especial 35º (750 cc)	***r95:*** Pisco especial 35º (645 cc)	***r96:*** Gasoline 93 octanes (lt)
r97: Gasoline 95 octanes (lt)	***r98:*** Gasoline 97 octanes (lt)	

Macroeconomic National Factors.

The macroeconomic national factors were obtained by principal components from the following variables:

- Source: EcoWin. Production, Manufacturing, Index, 2002=100 (Ew:clp02005); Labour Cost, Real, total, Constant Prices, Index , 2006M1=100 (ew:clp10020); Inactivity, Economic inactive population, total (ew:clp09030); Chile, Money supply M3, CLP (ew:clp12005); Light Crude Futures 33-Pos, Nymex, Close (ew:com2431510); OPEC Reference Basket Price, Average (ew:com2121010).
- Source: Central Bank of Chile. Interbank loan rate (1 day); Exchange rate from the central bank of Chile.

Table 1: Percentage of commodities for which each variable is significant (5%)

	Model M1	Model M2	Model M3	Model M4	Model M5	Model M6
Lagged dependent variable	100%	100%	100%	98%	95%	99%
Neighbour	-	81%	62%	76%	53%	71%
Lagged neighbour	-	32%	30%	13%	44%	7%
Common factor	41%	10%	91%	89%	67%	89%
Lagged Common factor	40%	9%	75%	58%	59%	41%
Aggregate Chilean inflation rate	-	23%	3%	5%	29%	3%
Lagged aggregate inflation rate	-	16%	2%	4%	17%	2%

Note: Model M1 – baseline model (Beck et al 2009 common factor); Model M2 – augmented model considering neighbours; Model M3 – national common factor for each commodity; Model M4 – M3 with contiguity matrix based on distance; Model M5 – GMM estimation of M4; Model M6 – M4 with variables in first differences.

Table 2: Average share of inflation variance explained by each variable

	Model M1	Model M2	Model M3	Model M4	Model M5	Model M6
Lagged dependent variable	45%	45%	45%	45%	45%	8%
Neighbour	-	7%	7%	7%	25%	0.2%
Lagged neighbour	-	21%	21%	23%	23%	1%
Common factor	16%	16%	31%	31%	31%	9%
Lagged Common factor	16%	16%	26%	26%	26%	1%
Aggregate Chilean inflation rate	-	15%	15%	15%	15%	1%
Lagged aggregate inflation rate	-	15%	15%	15%	15%	1%
Whole specification	46%	48%	52%	52%	47%	18%

Note: Model M1 – baseline model (Beck et al 2009 common factor); Model M2 – augmented model considering neighbours; Model M3 – national common factor for each commodity; Model M4 – M3 with contiguity matrix based on distance; Model M5 – GMM estimation of M4; Model M6 – M4 with variables in first differences.

Figure 1. Chile