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Informal Trade and the Price of Import Bans: Evidence from Nigeria

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Abstract

This paper studies the impact of trade restrictions and illegal trade on price disparities within Nigeria, where a long list of import prohibitions on various consumer goods have raised the price of the consumption basket. Large-scale illegal trade at the border with Benin allows traders bypass these prohibitions. We investigate whether this illegal trade mitigates the impact of import bans on domestic prices. Using a simple theoretical model and monthly price data, we exploit changes in the prohibition list, as well as spatial variation between border and non-border states, to identify the effect of smuggling. Results indicate that, for agricultural goods in particular, the price hike associated with an import ban is lower in border states. However, we find little evidence that bans disrupt the cointegration relationship between states. These results are consistent with a large diffusion of smuggled goods across Nigerian states, and with the higher price elasticity to distance in the illegal trade sector.

Résumé

Cet article étudie l'impact de la politique commerciale et du commerce illégal sur les écarts spatiaux de prix au Nigéria. Dans ce pays, de nombreuses interdictions d'importer s'appliquent à des biens de consommation variés. Pour contourner ces interdictions, un important commerce de contrebande avec le Bénin voisin s'est développé. Nous cherchons à déterminer si cette contrebande modifie l'impact des interdictions sur les prix domestiques. A l'aide de données de prix mensuels, nous exploitons les changements dans la liste d'interdictions, ainsi que la variation spatiale entre les états frontaliers du Bénin et les autres, pour identifier l'effet de la contrebande. Les résultats indiquent que le renchérissement lié à une interdiction d'importer est plus modéré dans les états frontaliers, surtout pour les biens agricoles. Cependant, les interdictions n'affectent pas la relation d'intégration entre les prix régionaux. Ces résultats sont cohérents avec une large diffusion des biens de contrebande dans l'ensemble des états nigérians, et avec des coûts de transport plus élevés dans le commerce illégal.

1. Introduction

Informal trade (i.e. unrecorded trade) is pervasive in Africa south of the Sahara and is often seen as the main explanation for low recorded levels of intra-African trade, despite the region's numerous integration schemes (Berg, 1985; Agbodji, 2007).

The impact of illegal trade or smuggling on welfare is ambiguous and remains a matter of debate in the literature[1]. Some authors argue that, by mitigating the impact of trade restrictions, smuggling may result in a higher welfare level than in a case of perfect enforcement of trade restrictions, in effect restoring part of the gains lost from trade (Deardorff and Stolper, 1990). The seminal paper by Bhagwati and Hansen (1973) reaches the opposite conclusion, however. In general, smuggling may entail a gain for consumers if it gives them access to cheaper goods relative to a case of perfect enforcement of trade restrictions; the size of these gains depends on the cost structure in both legal and illegal trade.

Nigeria has long imposed heavy trade restrictions, including import prohibitions and high tariffs. Import bans are intended to sustain domestic producers and develop local industries. However, the heavy cost of these bans on the purchasing power of the poor in particular, has been well documented (Treichel et al., 2010). This trade policy, combined with poor infrastructure (as well as restrictions on currency exchange), have created conditions for informal cross-border trade with neighboring countries, such as Benin, to develop on a large scale. It seems that, if there is a country where the case for welfare-improving smuggling can be made, it should be Nigeria. However, reaching this conclusion requires careful examination of the way in which local prices are formed, as well as of the interaction between the legal and illegal trade sectors. So far, a systematic study of the impact of import bans on prices across states within Nigeria, and of the interaction between bans and smuggling in the formation of prices, has not been undertaken.

This paper makes two contributions. First, it uses changes in the list of import prohibitions implemented between 2001-2010 to study the impact of these prohibitions on prices within Nigerian states. In particular, we test the hypothesis that the impact on state-level prices varies spatially within Nigeria; because of the importance of smuggling, states close to borders with neighboring Benin (a well-known corridor for smuggling) are likely to witness a weaker price response to import bans. Second, we test how import prohibitions modify integration between local markets within Nigeria, using cointegration analysis.

Our results are as follows. First, we see a clear positive impact of import bans on prices for a list of agricultural products, as well as an increase in price disparity across states. Bans increase the price of agricultural goods, with large variation across goods. Evidence of a similar impact for non-agricultural goods is less clear, possibly due to

[1] As explained in Cantens (2012), the notion of informal trade encompasses all unrecorded trade flows, this includes trade intended to avoid payment of taxes or bypass regulations (smuggling), as well as unreported local exchange within networks spreading across borders. In this paper we deal specifically with trade intended to bypass Nigerian bans and duties, and therefore use the terms illegal trade or smuggling.

non-homogeneity of goods. Second, for some of the products considered, there is evidence of a significantly smaller price hike in states bordering Benin relative to other states. The impact of bans is reduced by about one-third to one-half in these border states, suggesting that, for some goods, smuggling mitigates the impact of import bans in localities close to the border.

Third, our cointegration analysis indicates that state markets within Nigeria are integrated for goods under ban. Only in few cases do we find evidence that bans disrupt market integration. Considering prices in states bordering Benin and those in states distant from that border, we find that, in most cases, prices series remain linked by a long-term equilibrium relationship after a ban is imposed.

A small theoretical model is used to interpret these results; our findings are consistent with a level of trade cost such that importing goods illegally remains profitable, for the goods considered, in most states, including those distant from borders and ports. In other words, import bans do not choke off trade. The model shows conditions under which this holds true and shows that, if illegal trade entails higher transport costs than legal trade, one expects the price impact of bans to be lower in states closer to the border relative to more distant states.

The remainder of the paper is organized as follows. Section 2 gives an overview of the trade policy context in Nigeria and presents the data used. Section 3 presents the theoretical model used to clarify the mechanisms. Section 4 presents the empirical analysis. The last section concludes.

2. Context and Data

2.1 Informal Trade in Nigeria

Nigeria has long had a restrictive trade policy, with high tariffs and a long list of products prohibited for import. Along with poor infrastructure, in particular ports and roads, and currency controls, these policies have created an environment favorable for smuggling. Benin and Togo, two small countries neighboring Nigeria, have specialized as hubs for cross-border trade to Nigeria, so-called entrepot trade (Igue and Soule, 1992). There is abundant evidence regarding smuggling from Benin to Nigeria, including case studies (LARES, 2005) and estimates of informal flows using mirror trade statistics and other data (Raballand and Mjekiqi, 2010).

At the end of the 1990s, the government of Nigeria had a plan to phase out import bans. Only 23 products faced bans in 1998, and a number of these bans were replaced by high tariffs between 1999 and 2001. However, the trend reversed in 2002 (WTO), and a total of 218 products (HS 4-digit codes) faced an import ban at the end of 2004. These bans were imposed with the objective of protecting domestic industries and were maintained despite Nigeria's WTO membership, which should make such measures illegal, and despite the costs to the population and the loss of government revenue generated by such measures. According to close observers of domestic policy, Nigeria's government was well aware of these costs but was nonetheless determined to maintain the prohibitions for internal political reasons (Wikileaks, 2004).

These prohibitions remained essentially unchanged until the end of 2008, when a number of products were removed from the list; this was in part a consequence of Nigeria's membership of ECOWAS, which required it to align with the group's common external tariff. However, numerous products remained on the prohibition list after 2008, in a context in which the implementation of the ECOWAS common tariff was subject to negotiations and was repeatedly delayed.

2.2 Data

In order to measure the impact of import bans and protection measures on prices in Nigeria, we use three main sources of data. First, monthly time series for state-level retail prices in Nigeria were obtained from the National Bureau of Statistics (NBS). These series are collected to serve as input for the computation of consumer price indices; a list of about 150 products is included, which covers food, household goods, housing, clothing, and equipment. The list of products remained essentially unchanged over the period 2001-2010. Data for 2001-2006 provide one monthly price per state (there are 38 states in Nigeria). For 2007-2010, we include one urban and one rural price series.

We combine these price data with data on Nigerian trade policy measures, for which several sources are used. We combine prohibition lists as published by Nigeria's Minister of Finance with WTO reports to reconstruct a nearly complete history of import bans for the period 2004-2015. We also rely on secondary sources, such as Golub (2012) and Soule (2004), to check consistency and fill any gaps. We also use WTO and budgetary office documents for data regarding Nigeria's tariffs over the study period.

Finally, we also rely on the ECENE survey, a survey of informal border crossing points conducted by Benin's INSAE in 2010 and 2011, for data regarding informal trade flows between Benin and Nigeria. This survey collected data for about 10,000 informal trade flows across Benin's borders, a large majority of them at borders with Nigeria. This allows us to identify products for which informal trade exists; this is important because we expect the price impact of bans to differ depending on actual enforcement of import bans and on the persistence of trade.

This methodology allows us to determine that smuggling is indeed pervasive for products under ban in Nigeria. Among the 125 items listed in the NBS price data series, 47 faced a ban in 2011. Among these, 25 were identified in the ECENE database, meaning that informal flows were effectively recorded for these goods.

2.3 Descriptive Statistics

Table 1 displays descriptive statistics regarding the main variables used in the empirical analysis. Monthly state-level price data from the NBS are used for a range of consumer goods. Data for trade policy measures (ad-valorem data and prohibitions) at the national level are also used, with import bans coded as a binary variable. Table 2 displays detailed statistics regarding price data for agricultural products, on which part of the empirical analysis focuses. Overall, prices are relatively homogenous across states.

Table 11 in the Appendix lists the products that we use in the empirical exercise. These are all products present in the NBS price data for which there was at least one change in import ban status during our period of interest, 2001-2010; i.e. the product was included in the import prohibition list, or removed from it, during that period. Our empirical specification will rely on such changes to identify the impact of bans on prices. In practice, one observes that there were two main changes in the list during the study period: a number of products were added in 2003 and 2004 and another group was excluded in October 2008 (noted 2009).

Table 1: Descriptive statistics

Variable	Mean	SD	Min	Max	Nb. Obs
Log price	4.95	1.19	-3.2	9.15	124363
Import ban	0.77	0.42	0	1	124363
Log (1+tariff)	0.22	0.14	0.049	0.69	124363
Ecene	0.49	0.5	0	1	124363

Sample: 2001-2010 state-level monthly prices. Source: NBS, WTO, INSAE. Ecene: 1 if product appears in ECENE survey on informal trade at Benin-Nigeria border crossing points.

Table 2: Descriptive statistics: agricultural prices (Naira)

	Mean	Std. error	Min	Max	Nb. Obs.
			Chicken (unit)		
Jan. 2001	342.9	70.6	120	448	
Dec. 2010	1310.4	194.9	1000	2066.6	37
2001-2010	759.7	308.1	120	2066.6	4404
			Plantain (1 kg)		
Jan. 2001	47.8	14.4	24.3	74.7	35
Dec. 2010	184.2	55.7	150	333.3	36
2001-2010	114.2	67.8	24.3	935.5	3964
			Gari (white, 1 kg)		
Jan. 2001	43.2	13.1	21.7	67.7	36
Dec. 2010	113.3	18.8	65.2	177.8	37
2001-2010	68.9	22.7	21.7	204.1	4428
			Guinea corn (1 kg)		
Jan. 2001	27.5	8.6	14.1	46.9	36
Dec. 2010	80.1	28.3	41.4	141.7	37
2001-2010	57.3	26.6	14.1	166.8	4392
			Maize grain (white)(1 kg)		
Jan. 2001	28.9	8.4	16.8	46.2	36
Dec. 2010	70.6	20.6	42.6	111.1	37
2001-2010	54.7	23.6	15.3	164.7	4404
			Millet (1 kg)		
Jan. 2001	29.6	9.7	14.4	46.6	36
Dec. 2010	77.2	23.8	36.5	125	37
2001-2010	82.7	85.1	14.4	400	4392
			Vegetable oil (1 litre)		
Jan. 2001	104.3	8.5	79.7	125	36
Dec. 2010	272.7	27.1	217.3	400.1	36
2001-2010	193.9	49.7	79.7	422.2	4425
			Palm oil (1 litre)		
Jan. 2001	81.7	8.3	67.4	106.3	36
Dec. 2010	254.3	10.3	231.6	271.1	36
2001-2010	169.8	49.7	57.1	300	4425

Monthly, state-level prices in Naira from NBS data.

3. Theoretical Model

Theoretically, the impact of import bans on prices depends on the costs of trading goods both legally and illegally, as well as on the prices of imported and domestic goods. Depending on these costs, which vary across states, it may or may not be profitable to import goods from abroad. In order to clarify these mechanisms, we build a small model for trade in Nigeria, a country which imposes tariffs and import bans on some goods. Thus, in Nigeria, there exists the possibility of smuggling goods from abroad, at costs which we will specify. We are interested in knowing which Nigerian states will import from abroad versus which will consume local production, and how are prices impacted in each scenario.

Let us define as $p^{W}{}_{it}$ the world price of good i at date t on the world market: for now, we will make a small open economy hypothesis and assume that this price is exogenous to market conditions in Nigeria. We define this price as the price at the port in Nigeria, including international transport costs but excluding tariffs. This is the CIF (cost insurance freight) price of imports. In our model, we will be interested only in internal costs and their impact on local prices; we do not model international trade costs.

We note τ_s as the internal transport cost of legal trade: the cost of transporting goods legally from the border (e.g. the port of Lagos) to state s within Nigeria. This cost is modeled as equivalent to an ad-valorem tax on the value of the shipment excluding tariffs: i.e. this cost is equal to $\tau_s.p^{W}{}_i$. For now, we assume it to be constant across goods.

In the period considered (2001-2010), all of Nigeria's tariffs are ad-valorem (no specific tariffs). We note T_i as the ad-valorem tariff on good i.

Let us consider the price of a good i in state s. We first assume that there is no smuggling. If it is profitable to import the good legally into state s, then the price of good i in state s is given by

$(1+T_i+\tau_s).p^{W}{}_i$. If importing is too costly relative to local production, then $p^{NT}{}_{is}$ denotes the price of the good in state s without trade. Therefore the price can be written as:

$$p_{is} = \min[(1 + T_i + \tau_s).p_i^W, p_{is}^{NT}] \qquad (1)$$

We now consider the possibility of smuggling. This trading mode entails specific costs. First, the transport cost for smuggling is modeled as:

$$\tau_s^{Smg} = (1 + \alpha_s)\tau_s \qquad (2)$$

This is intended to include the additional costs associated with alternative transport routes which have to be used by smugglers, relative to the legal trade route. These costs may take the form of delays, as well as the damage or possible loss of a share of goods transported this way. As for legal trade, we assume these costs apply to the CIF value of the good, excluding tariffs (i.e. the costs are equal to $(1 + \alpha_s)\tau_s.p^{W}{}_i$).

Second, smuggling requires the payment of an informal tax or bribe. In the case of goods facing tariffs but no ban, we write this tax as

$$IFT_i^{Smg} = (1 - \gamma_i)T_i \qquad (3)$$

with $0 < \gamma_i < 1$. Goods smuggled into the country pay no tariff but do pay this tax instead; therefore their price at the border, before internal transport, is equal to $p_i^W.(1 + (1 - \gamma_i)T_i)$. In the case of a product under ban, we assume the informal tax is at IFT_i^B.

Finally we model repression of smuggling as entailing an additional cost on trade. We note ψ_s^B and ψ_s^{NB} the ad-valorem equivalents of this cost under the ban and no ban regimes, respectively, which depend on the risk of being arrested. It can be considered as a premium paid on delivered goods serving as an insurance payment against this risk. We model these costs as applying to the CIF value of the goods (i.e. $p^{W}{}_{i}$). We assume that the smuggling sector, as the legal distribution sector, is perfectly competitive.

Summing up, the price of a good *i* in state *s* will be equal to:

- $(1 + T_i + \tau_s)p_i^W$ *if the good is imported legally into state s*
- $\left[1 + (1 + \alpha_s)\tau_s + (1 - \gamma_i)T_i + \psi_s^{NB}\right] p_i^W$ *if the good is smuggled into state s, in the case of a product not under ban*
- $\left[1 + (1 + \alpha_s)\tau_s + IFT_i^B + \psi_s^B\right] p_i^W$

if smuggled into state*s*, in the case of a product facing a ban.

Note that we will have to ask whether and when such imports will be less costly than local production; only if they are less costly will these prices represent the actual price on the market in state *s*.

We first consider a state *s* with low-cost access to the world market (low τ_s), so that it is always profitable to import good *i* from abroad. If tariffs are not too high, then in the absence of an import ban, goods will be imported legally. If a ban is imposed, then goods will be imported through the smuggling route. Thus, the price gap due to the ban can here be written as:

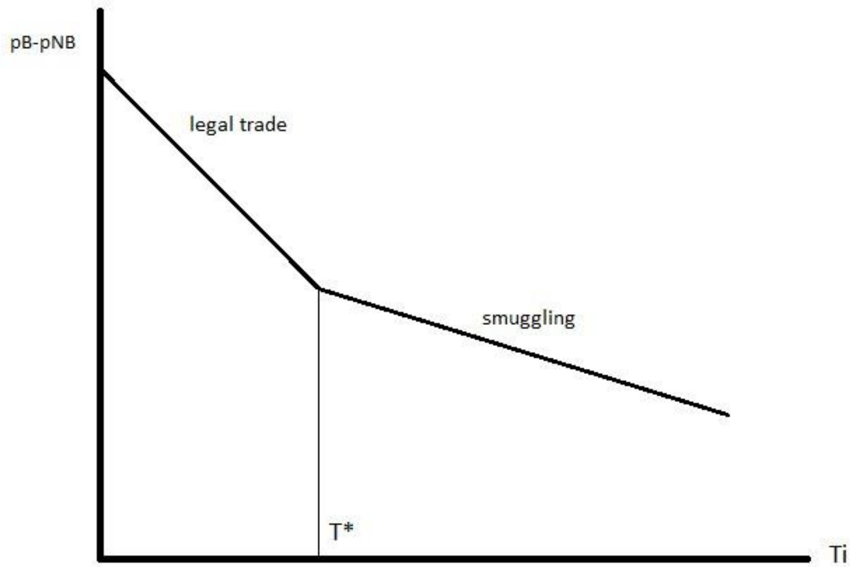
$$p_{is}^B - p_{is}^{NB} = [IFT_i^B + \alpha_s\tau_s + \psi_s^B - T_i].p_i^W \tag{4}$$

This will hold provided that, in the case of no import ban, legal trade is indeed cheaper than smuggling; this is true as long as $\alpha_s\tau_s + \psi_s^{NB} > \gamma_i T_i$. When this inequality does not hold, then the good is being smuggled both under the ban and the no ban regimes. The price difference then writes:

$$p_{is}^B - p_{is}^{NB} = [IFT_i^B - (1 - \gamma_i)T_i + \psi_s^B - \psi_s^{NB}].p_i^W \tag{5}$$

The relationship between the price gap $p^B{}_{is} - p^{NB}{}_{is}$ and the tariff is illustrated in Figure 1. This gap decreases with the tariff, such that a ban has less impact on the price of a good which was already facing a high tariff. However, the relation is non-linear, with the slope being steeper in the low-tariff zone in which the tariff increases the price under no ban but has no impact on the price in the ban regime (equation 4). There is a threshold T^* at which trade becomes illegal when there is no ban; after this threshold, the slope of the gap is less steep, as the tariff also impacts the legal price, through the informal tax, or bribe, which has to be paid.

Figure 1: Price impact of a ban as a function of the tariff



We now consider this price gap as a function of the transport cost, τ_s. Under no ban, and assuming that tariffs are not too high, the price will be the lowest between the local, no-trade price, and the price of legal imports:

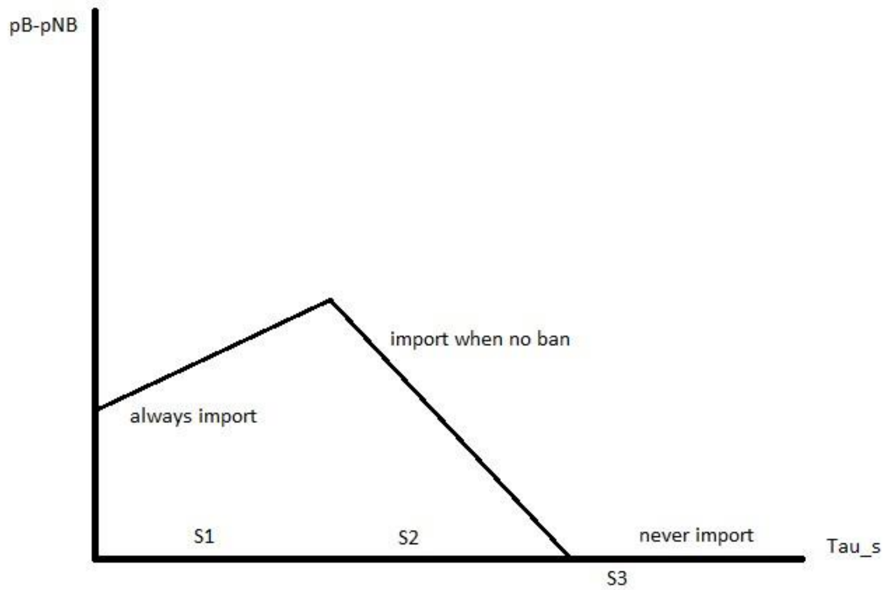
$$p^{NB}_i = min[(1 + T_i + \tau_s)p^W{}_i ,p^{NT}{}_i] \quad (6)$$

Under the ban, goods will be smuggled, provided that this remains cheaper than local production:

$$p_i^B = \min\left\{\left[1+\psi^B+IFT^B+(1+\alpha_s)\tau_s\right]p_i^W, p_i^{NT}\right\} \quad (7)$$

The resulting price gap is represented in Figure 2. Is the price impact of a ban larger in more remote states than in those closer to the world market? This figure shows that the answer is not clear. If we know that smuggling is always profitable in the places considered (such as state S1 on the graph), then the price gap increases with distance (trade cost), as seen in the first segment of the graph. Note that the slope is α_s; thus if there is no trade cost component specific to smuggling (e.g. smugglers follow normal routes and only pay an informal tax), then the price difference is flat: close and remote places all pay the same premium ($IFT + \psi^B - \psi^{NB}$) for having goods delivered.

Figure 2: Price impact of a ban as a function of transport costs



However, if one compares a state (such as S1 in the figure) in which smuggling occurs with a state in which the ban chokes off trade (S2), then the answer is ambiguous: the price hike may be higher in either state.

This analysis shows that the price impact of a ban depends on whether trade continues (illegally) after the ban or not. One method to test for this will be to test for market integration between states. We expect that markets in which goods are smuggled exhibit integration (i.e. cointegration between prices), as prices remain anchored to the world price. By contrast, an absence of integration between prices should indicate that international trade has been choked off at least in some of the states considered.

Empirical specification. Consider states in which goods are always imported: legally under no ban and illegally if a ban is imposed (case S1 in Figure 2.) Then the price p_{ist} in state *s* can be written as:

$$p_{ist} = p_{it}^{W} \cdot \left[(1 + T_i + \tau_s) + Ban_{it}.(\alpha_s \tau_s - \gamma_i T_i + \psi^{NB})\right]. \qquad (8)$$

where Ban_{it} is a binary indicator for an import ban on good *i*. Log-linearizing (assuming small tariffs and trade costs) , one obtains:

$$\ln p_{ist} = \ln p_{it}^{W} + T_i + \tau_s + (\alpha_s \tau_s - \gamma_i T_i + \psi^{NB}).Ban_{it} \qquad (9)$$

This equation will serve as the basis for our econometric specification; we will use product-time fixed effects, which will control for variations in the world price of the good. State fixed effects will capture differences in transport costs for each state. We will include a binary variable Ban_{it} and interact it with tariffs T_i; the coefficient on the interaction is expected to be negative. Finally, we will also interact Ban_{it} with an indicator for border states, i.e. states bordering Benin. We expect trade costs (τ_s), and in particular the specific trade costs associated with smuggling (α_s), to be lower in these states. Therefore, we will expect to find a lower impact of bans in these states (i.e. a negative coefficient on the interaction Ban * Border).

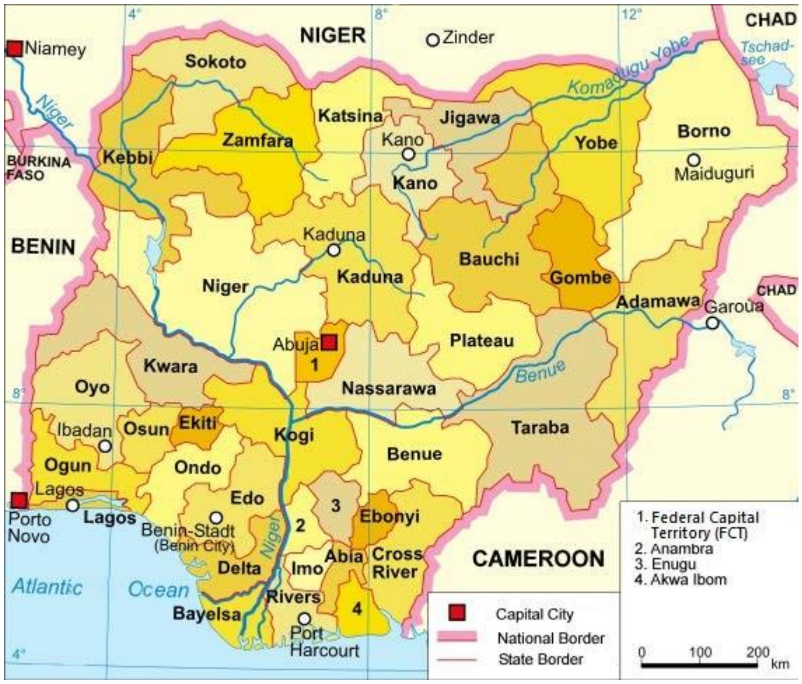
Note that, as the theoretical exercise has shown, the sign of this interaction depends on whether states continue importing under the ban. Therefore, we will use a cointegration analysis to test whether the prices of goods under a ban exhibit equilibrium relationships across states.

4. Impact of Trade Restrictions on Prices in Nigeria

We ask how restrictions on imports, in particular import bans, impact local prices in Nigeria. We expect these restrictions to raise the price of the goods facing the restriction by an amount which depends on the nature of the restriction (whether a ban or a tariff) and on demand and supply elasticities, as well as on the enforcement of the restriction. One hypothesis is that price impacts may vary spatially, if states close to borders may receive flows of smuggled goods which mitigate the restriction.

A political map of Nigeria is displayed in Figure 3. There are 36 states on this map, which will be our spatial unit of observation for prices. Six states border Benin: Lagos, Ogun, Oyo, Kwara, Niger, and Kebbi. There is abundant evidence of informal trade across the border between Benin and Nigeria (Igue and Soule, 1992; Golub, 2012). Therefore, we expect the price impact of trade restrictions to potentially differ in these states relative to other states.

Figure 3: Nigeria States



4.1 Graphic Evidence

We start by looking at a few examples of products that have faced import bans in Nigeria. We focus in particular on cases in which a ban was either imposed or lifted during the period for which we have price data.

Figures 4 and 5 display price series for palm oil and vegetal (other than palm) oil in a selection of states in Nigeria. A ban was imposed on imports of vegetable oil in bulk in September 2003 (WTO, 2008). We consider six states of Nigeria. Kebbi, Ogun, and Lagos share a border with Benin, while Abuja, Plateau and Gombe are further away from the border and are also at distance from other borders (see the map in Figure 3).

The impact of the ban on the price of vegetable oil is clear in Figure 4. The ban appears to increase price dispersion across states, and may also increase price volatility. Prices appear to decrease in Kebbi and Ogun, two states bordering Benin.

The impact is less clear in the case of palm oil, but the figure also suggests an increase in price dispersion. It is possible that the prohibition was implemented earlier, e.g. in late 2002, as we are not certain about the precise date. Price volatility seems higher for all states in the post-ban period.

We conduct a similar exercise for the price of chicken. A ban on imported poultry was put in place in 2004 (after having been lifted after 1997). Figure 6 shows that the price of chicken rose strikingly in January 2004. We observe this rise in all states but Oyo. We use Oyo instead of Kebbi for this product because prices in Kebbi appear constant, and are probably inaccurate. Again, we observe price dispersion to increase following the ban. This suggests that regional markets for chicken may have been integrated in the pre-ban period, but were not integrated after 2004.

One explanation for this could be that smuggled goods reach some states but not others; these latter states then become reliant on local production. Indeed, one also observes that, in the post-2004 period, prices increase more steeply in Abuja and Gombe than elsewhere. This could be related to the geographical position of these states, which are distant from the border with Benin. Whether this type of relationship can be verified in a systematic manner is the question we will address in the next section, using difference-in-difference analysis.

In Figure 7, we look at the price of local chicken: how is that market impacted by the import ban? Supposedly, local chickens are substitutes for imported ones. One observes that, even before the ban, price dispersion is higher for local than for imported chicken, presumably reflecting regional differences in production costs and internal transport. The ban has an impact on this market, particularly in Abuja where the price hike is clear. Again, price dispersion increases post-ban, and the gap between border and non-border states appears to increase.

Turning to the case of maize (seen in Figure 8), we see that this staple was included in the prohibition list from January 2004 to October 2008. The price of maize started to rise in 2007-2008, likely due to the upward trend in world food markets in that period. This example shows the difficulty in identifying the impact of bans which were removed in 2008. It seems that this removal contributed to putting an end to the upward price trend, but disentangling the effects of ban removal and of world market fluctuations is difficult in that period.

This is also the case to some extent for millet, shown in Figure 9, although prices seem to have clearly decreased after October 2008 in the states of Kebbi, Gombe, and Plateau.

The case of millet also provides an interesting example of price disparities across states. Nigeria is a large millet producer, and local prices likely reflect production costs and internal transportation. The price is generally higher in Lagos, probably because of costs of transport to this urban center. Notably, the Lagos premium shrank during price increases in mid-2005 and in 2008, suggesting asymmetry in price adjustments to price fluctuations and the possibility that this premium is in part due to margins in the distribution sector.

Overall, these figures show that, at least in some cases, there is a clear positive impact of import prohibitions on prices, that these prohibitions tend to increase price disparities across states, and possibly that the price impact of bans is mitigated in states bordering Benin relative to other states. We also note that some of the price series are quite noisy, which will make identification of the impact of bans more difficult, as does the absence of accurate information regarding the precise date of implementation of import restrictions in many cases. Finally, we note that for some of the products considered, a graphic display suggests that price volatility may rise following the implementation of a ban.

Figure 4: Price of Vegetal oil, by state

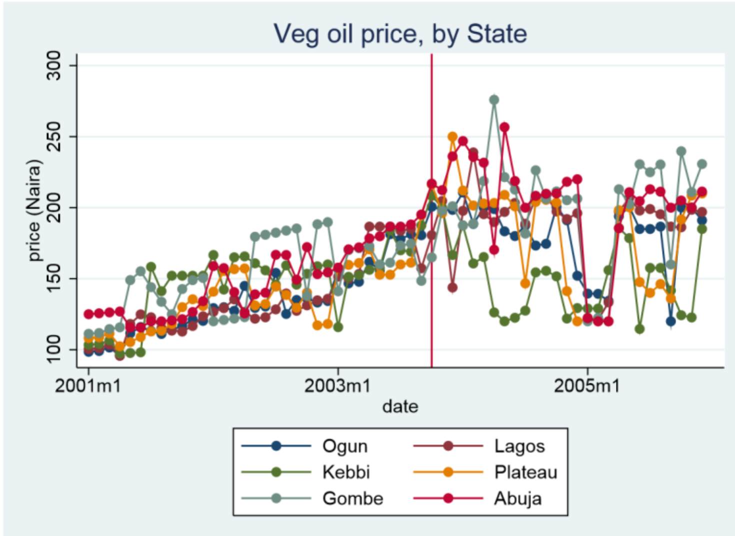


Figure 5: Price of palm oil, by state

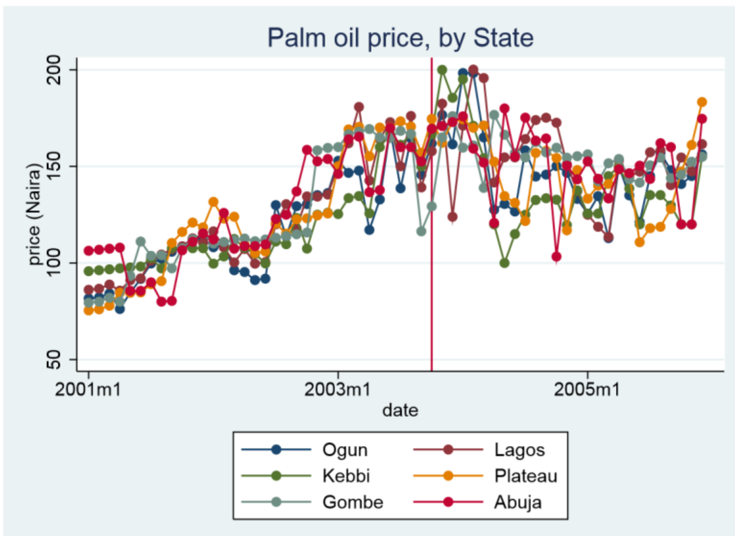


Figure 6: Price of Chicken, by state

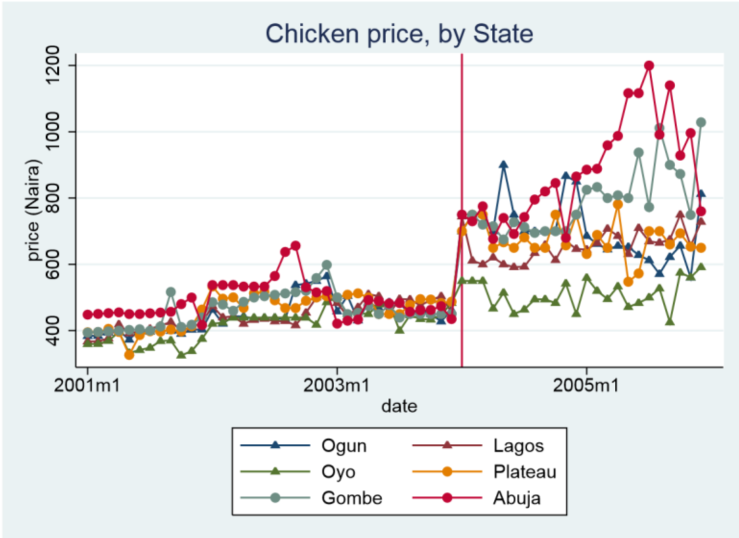


Figure 7: Price of local chicken, by state

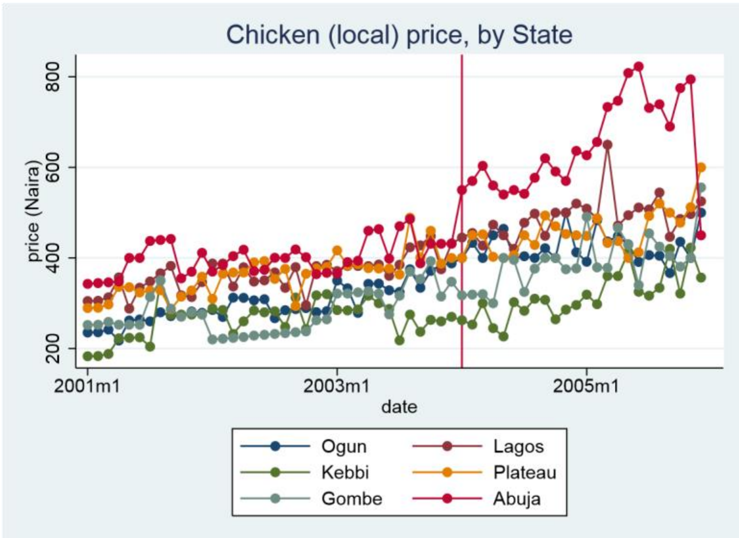


Figure 8: Price of Maize, by state

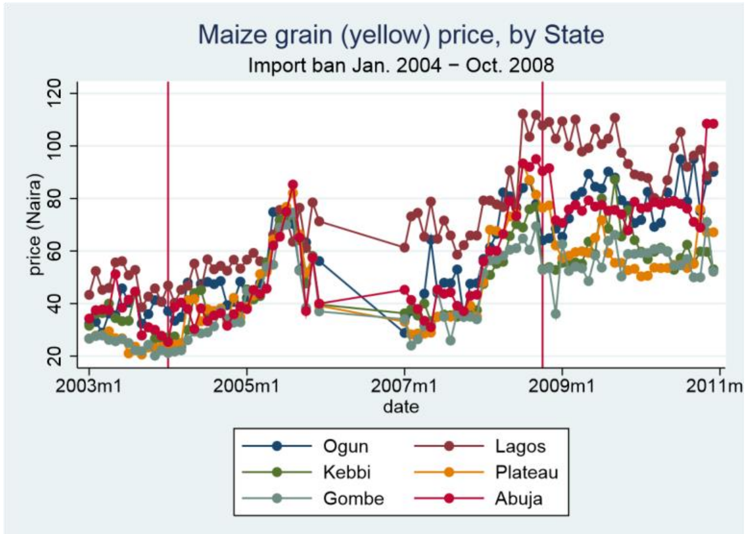
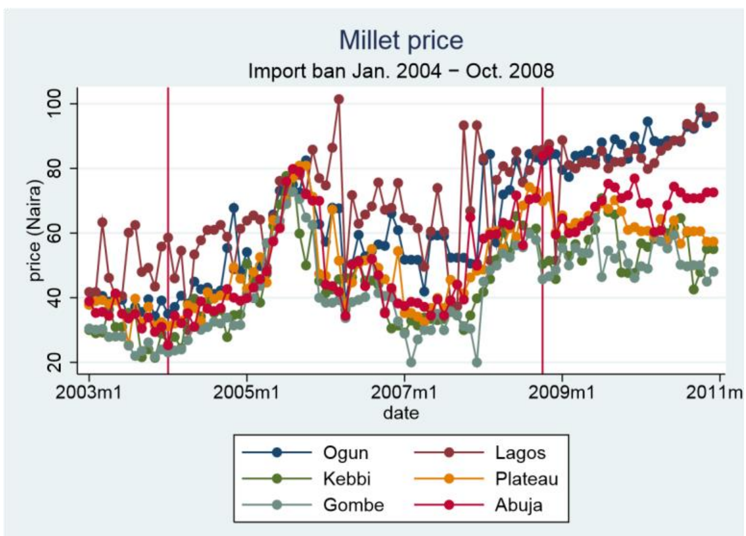


Figure 9: Price of Millet, by state



4.2 Stationarity

In order to test for stationarity of the price series, we use an augmented Dickey-Fuller test on GLS detrended data for monthly price series for each state and product, allowing for up to 12 lags. We select the optimal number of lags using the Ng-Perron criterion. Structural breaks are likely to be present in the data, biasing the results toward non-stationarity (Perron and Vogelsang, 1992a). One cause for such breaks could be changes in ban status. We address this issue first by running the tests on periods when the ban was constant. We focus on products for which the ban status changed once during the period, split the sample in the pre-ban and post-ban periods (2004-2008 or

2004-2010 in most cases), and run the test on these two periods. Results are displayed in Table 3, showing that the null hypothesis of a unit root is rejected in a number of cases.

Another way to address the issue of structural breaks is by using Perron and Vogelsang's (Perron and Vogelsang, 1992b) test, which allows for one structural break. In the additive outlier model, the procedure selects the timing of the break and then tests the hypothesis of a unit root in a model with a shift in the level of the variable at the breakpoint.

These tests indicate that non-stationarity is rejected in a number of cases (states). There is no clear indication that prices would be uniformly I(1), possibly because of market imperfections. This is further confirmed when using panel unit-root testing; for each product listed in Table 3, we test the hypothesis that all State prices have a unit root using Im Pesharan and Shin's test: the null hypothesis is rejected at 1% for every product. Therefore, the tests indicate that at least some of the series are stationary.

Non-stationarity creates a risk of spurious regressions in OLS, and our primary methodology should therefore rely on a cointegration analysis. In the next section, however, we will test the impact of bans using level regressions to estimate the change in price level, focusing in particular on products for which prices appear stationary. Then, in Section 4.4, we will analyze cointegration relationships between state prices.

Table 3: Stationarity: Dickey-Fuller GLS tests on constant ban periods

Item	**Nb. H0 rejected (5%)**		**Nb. series**
	Period 1	**Period 2**	
Banana	5	27	36
Chicken-agric(medium)	4	31	36
Gari-white	12	1	36
Gari-yellow	15	10	36
Guineacorn	6	3	36
Maize-grain(white)	5	6	36
Millet	6	0	36
Orange	7	29	36
Palmoil	0	19	36
Plantain-ripe	8	23	36
Vegetableoil	4	34	36

Table 4: Stationarity: Perron Vogelsang tests

Item	Nb. H0 rejected	Nb. series
Banana	14	36
Chicken-agric(medium)	10	36
Gari-white	19	36
Guineacorn	21	36
Maize-grain(white)	23	36
Millet	4	36
Palmoil	13	36

4.3 Difference-in-Difference Analysis

We employ a difference-in-difference methodology to investigate the impact of import bans on price; we will exploit time variation in protection, focusing on cases in which a ban was imposed or lifted within the period covered by our price data. We will also use the comparison between border and non-border states; our prior is that access to smuggled goods may be easier in border states, so that the impact of a ban on prices is not uniform across state. Our baseline specification will be the following:

$$\ln p_{ist} = \alpha_{is} + \lambda_i * t + \beta_1.BAN_{it} + \beta_2.\ln(1+AV_{it}) + \beta_3.BAN_{it}.\ln(1+AV_{it}) + \gamma.BAN_{it} * BORDER_s + \epsilon_{ist} \tag{10}$$

where $\ln p_{ist}$ is the log of the price of product i in Nigeria's state s at the (monthly) date t. α_{is} are product-state fixed effects; $\lambda_i * t$ is a product-specific linear trend (with t indicating date in months).

BAN_{it} is a binary variable indicating that product i faces an import ban at date t. $\ln(1 + AV_{it})$ denotes the log of the ad-valorem tariff applied on imports of product i at time t. The model includes an interaction term between the Ban and tariff variables, designed to control for the interaction between the two forms of import restrictions.

Finally $BAN_{it} * BORDER_s$ is an interaction term between the ban variable and an indicator for border states (states with a border with Benin). $_{ist}$ is an error term supposed to be normally distributed.

Fixed effects α_{is} are used to control for the permanent component of differences across states in the prices of goods. Such differences may be due to states' position relative to places of production and distribution of the goods (e.g. location of production of goods in the country, distance to ports and borders, and transport costs). In addition, the term $\lambda_i * t$ controls for changes in product prices that are common to all states (e.g. world price changes).

Tables 5 and 6 display results from estimations of the baseline specification. In Table 5, we test the impact of import bans and tariffs on prices; in Table 6, we allow for spatial variation in this impact.

Import bans generally have the expected positive impact on prices, with an average impact of about 5 percent when excluding medicaments, for which the impact is negative. As expected, ad-valorem tariffs increase retail prices, and the interaction with bans is negative; this interaction must be controlled for, as import tariffs are frequently modified along with bans.

In column 3, we include a proxy for smuggling. We expect the price impact of bans to be smaller when a banned product can be smuggled. We use data from the 2011 ECENE survey on informal trade in Benin to identify the list of products appearing in informal flows from Benin to Nigeria. This provides an indicator of smuggling, given that ban lists did not change completely until between 2003 and 2011 and that we expect some persistence in smuggling networks. However, this variable attracts a negative but small coefficient, possibly due to measurement error.

Considering product categories, the impact is strongest in agricultural goods. We also detect an impact in at least some manufactured goods.

Table 6 introduces the interaction between import bans and border states (the six states bordering Benin). Overall, the impact of the ban is reduced by half on average in border states relative to other states; however, the effect is only significantly detected for agricultural products.

In Table 7, we focus on agricultural goods. Product-specific regressions identify larger impacts. This effect is reduced by one-third to one-half in border states.[2,3]

Table 5: Impact of import bans on prices

	Dep. variable: lnp_{ist}							
		no meds.		**Agro**	**Agro-Indus.**	**Meds.**	**Textile.&App.**	**Other Manuf.**
Import ban	0.031^{a}	0.052^{a}	0.056^{a}	0.060^{a}	-0.317^{a}	-0.034^{a}	0.021	0.024^{b}
	(0.004)	(0.004)	(0.005)	(0.005)	(0.030)	(0.004)	(0.019)	(0.011)
Ad-valorem tariff	0.025^{b}	0.076^{a}	0.075^{a}	0.010	-1.057^{a}		0.123	0.117^{a}
	(0.013)	(0.014)	(0.014)	(0.019)	(0.114)		(0.078)	(0.025)
Ban * tariff	-0.086^{a}	-0.119^{a}	-0.118^{a}	-0.117^{a}	1.121^{a}		-0.105	-0.126^{a}
	(0.011)	(0.012)	(0.012)	(0.016)	(0.112)		(0.074)	(0.024)
Ban * Ecene			-0.008^{c}					
			(0.004)					
Observations	124363	94639	94639	53125	10269	29724	7089	24156
R^2	0.533	0.550	0.550	0.565	0.639	0.465	0.217	0.531

Standard errors in parentheses.
^{c} $p < 0.10$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

[2] The negative coefficient in the case of Gari (Manioc our) may be explained by the fact that for this staple, import ban was removed in 2008, contrary to other products where we observe a ban being imposed in 2003 or 2004. Asymmetry in price reactions, in conjunction with high price volatility in the 2008-2009 period, are the most likely cause of this result.

[3] We also experimented with a specification including product-time (month) fixed effects, as well as state fixed effects, in order to better control for changes in e.g. world prices. Results for the interaction border states/ban are essentially unchanged.

Table 6: Impact of import bans on prices: border vs. non-border states

		Agro	**Agro-Indus**	**Meds.**	**Textile.& App.**	**Other Manuf.**
Import ban	0.033^{a}	0.065^{a}	-0.318^{a}	-0.031^{a}	0.025	0.021^{b}
	(0.004)	(0.005)	(0.030)	(0.005)	(0.019)	(0.011)
Ban * border State	-0.016^{a}	-0.029^{a}	0.017	-0.018^{c}	-0.025	0.015
	(0.004)	(0.007)	(0.012)	(0.009)	(0.016)	(0.009)
Ad-valorem tariff	0.025^{b}	0.010	-1.052^{a}		0.123	0.117^{a}
	(0.013)	(0.019)	(0.114)		(0.078)	(0.025)
Ban * tariff	-0.086^{a}	-0.117^{a}	1.116^{a}		-0.105	-0.126^{a}
Observations	124363	53125	10269	29724	7089	24156
R^2	0.533	0.565	0.640	0.465	0.217	0.531

Standard errors in parentheses.
^{c} $p < 0.10$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

Table 7: Impact of import bans on prices: by product

	Chicken	**Plantain**	**Gari**	**G.corn**	**Maize**	**Millet**	**Veg.oil**	**PalmOil**
Import ban	0.370a	0.319^{a}	-0.133^{a}	0.106^{a}	0.135^{a}	0.571^{a}	0.159^{a}	0.201^{a}
	(0.013)	(0.026)	(0.014)	(0.024)	(0.023)	(0.032)	(0.010)	(0.012)
Ban * border	-0.083^{a}	0.011	-0.087^{a}	0.042	-0.066^{b}	0.013	-0.008	-0.028^{c}
	(0.016)	(0.027)	(0.019)	(0.032)	(0.027)	(0.043)	(0.015)	(0.015)
trend	-0.003a	-0.003	0.015^{a}	0.016^{a}	0.009^{a}	0.003	0.004^{a}	0.004^{a}
	(0.001)	(0.004)	(0.001)	(0.003)	(0.002)	(0.003)	(0.000)	(0.000)
R^2	0.814	0.607	0.680	0.696	0.644	0.491	0.772	0.814
Observations	3540	1728	3552	2640	2640	3540	4425	4425

Standard errors in parentheses.
^{c} $p < 0.10$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

4.4 Market Integration

We now turn to the analysis of integration between state markets within Nigeria. Results in the former section show that import bans raise the price of goods; for some goods at least, we have evidence that this effect is lower in states bordering Benin, suggesting a role for smuggling from this country. These results suggest that smuggled goods may be more pervasive in border markets than in those more distant from the border.

As shown in the model in Section 3, under some parameter values, smuggling goods may remain profitable in markets close to the border, while more distant markets will be cut from the international market and will have to source goods locally. In this case, one would then expect to find a weakened relation of integration between such markets. This section tests this hypothesis.

Table 8 summarizes results from a series of Johansen tests for cointegration for a selection of goods facing an import ban and for a selection of border and non-border states. We focus on a list of agricultural goods produced regionally. (Note that these are also rather homogenous goods.) For each of these products, we test for a binary cointegration relationship between each of three states bordering Benin (Lagos, Ogun, and Oyo) and three states distant from Benin: Abia (Southeast), Plateau (Center-East) and Borno (Northeast, most distant from Benin; see the map of Nigerian states in Figure 3).

Johansen tests are conducted on pairs of price series (one border and one non-border state), focusing, for each good, on the period covered by the ban (2004-2010, 2004-2008, or 2003-2010 depending on the good). We test the hypothesis of a maximum rank of 0 using the trace statistic, at the 95 percent level of confidence. Lag selection based on the AIC criterion indicates in most cases an optimal lag number of 2 in the VAR model, corresponding to one lag in the VEC model. We thus include one lagged term in the VECM. [4]

Results show a fairly high degree of integration. For all six goods considered, we find evidence of cointegration between at least some of the markets considered. Integration is strongest for vegetable and palm oil.

An alternative approach to testing for long-term relationships has been proposed by Pesaran et al. (2001); this approach is suitable when the series' degree of integration is uncertain: (I(0) or I(1)), as is the case here. Using the bounds testing procedure proposed by these authors, we obtain the results summarized in Table 9. These results are broadly consistent with those in Table 8 and confirm that the prices considered exhibit a high level of integration overall.

Next, we test for structural breaks in the long-term equilibrium relationship between prices: do import bans modify this relationship? To investigate, we estimate a modified error-correction model of the form:

$$\Delta p_{st} = c + (\alpha + \alpha_1 . Ban_t).(p_{s,t-1} - \beta \ p_{s',t-1}) + \sum_{i \geq 1} \gamma_i \Delta p_{s,t-i} + \sum_{j \geq 0} \delta_j \Delta p_{s',t-j} + \epsilon_t$$

where p_{st}, p_{s0t} are monthly prices in state s and s^0, and Ban_t is a dummy for import ban (omitting the product index for clarity). A positive coefficient, $\alpha_{1,}$ indicates that adjustment toward equilibrium is slower when a ban is imposed (or that there is no adjustment).

Results confirm the presence of a ban-induced structural break only in a minority of cases. As an example, Table 10 displays results for vegetable oil prices. In four cases out of the nine pairs of markets considered, we find a significant reduction of adjustment under a ban. However, we do not find similar results for other products. One reason for this is the small number of data points, with generally fewer than 40 months of data in the period before the ban is imposed. This is probably too short a time period to properly identify a regime change.

[4] We test with two vector error-correcting models, with and without a constant in the specification. We reject the hypothesis of zero rank when it is rejected in both models, following Ahking (2002).

Our results in this section indicate that state markets for the goods considered are most often integrated across the country, even in periods of import bans. We find limited evidence that import bans weaken the equilibrium relationship between prices.

Table 8: Cointegration results (Johansen test): summary

Item	**period**	**Lagos**		**Ogun**			**Oyo**			
		Plateau	**Abia**	**Borno**	**Plateau**	**Abia**	**Borno**	**Plateau**	**Abia**	**Borno**
Chicken	2004-2010	C			C	C	C	C		
Gari	2003-2008		C	C		C	C		C	C
Maize	2004-2008				C		C	C		C
Veg oil	2003-2010	C	C	C	C	C	C	C	C	C
Palm oil	2003-2010	C	C	C	C	C	C	C	C	C
Millet	2004-2010			C			C			C

Summary of results of Johansen tests for cointegration between prices in states bordering Benin (Lagos, Ogun, and Oyo) and non-border states (Plateau, Abia, and Borno). C indicates that the hypothesis of zero maximum rank was rejected in pairwise tests (trace statistic), both with and without a restricted constant.

Table 9: PSS bounds testing procedure for long-term equilibrium relationships

Item	**period**	**Lagos**		**Ogun**			**Oyo**			
		Plateau	**Abia**	**Borno**	**Plateau**	**Abia**	**Borno**	**Plateau**	**Abia**	**Borno**
Chicken	2004-2010	LTE	LTE	LTE	LTE	LTE	LTE	LTE		
Gari	2003-2008	LTE			LTE	LTE	LTE	LTE		
Maize	2004-2008				LTE		LTE	LTE		LTE
Veg oil	2003-2010	LTE	LTE	LTE	LTE	LTE	LTE	LTE	LTE	LTE
Palm	2003-2010	LTE	LTE	LTE	LTE	LTE	LTE	LTE	LTE	LTE
Millet	2004-2010	LTE		LTE	LTE		LTE		LTE	LTE

Summary of results of Pesaran, Shin, and Smith (2001)'s bounds testing procedure for long-term relationship. LTE indicates rejection of H0: no long-term relationship at 5% (F-stat > critical value for I(1) regressors).

Table 10: Error correction model: vegetable oil prices

LHS price	Plateau			Abia			Borno		
RHS price	**Lagos**	**Ogun**	**Oyo**	**Lagos**	**Ogun**	**Oyo**	**Lagos**	**Ogun**	**Oyo**
Import ban	0.100	0.014	-0.003	0.008	0.070[a]	0.040	-1.332[a]	-0.955[a]	-0.915[a]
	(0.129)	(0.038)	(0.027)	(0.021)	(0.027)	(0.037)	(0.392)	(0.212)	(0.241)
Error correction	-0.252	-0.296	-0.267	-0.553[a]	-0.822[a]	-0.581[a]	-1.595[a]	-1.553[a]	-1.664[a]
	(0.240)	(0.265)	(0.264)	(0.187)	(0.163)	(0.182)	(0.322)	(0.228)	(0.288)
Ban *	-0.210	-0.207	-0.328	0.097	0.410[b]	0.137	1.126[a]	1.088[a]	1.122[a]
Error correction	(0.255)	(0.278)	(0.278)	(0.204)	(0.178)	(0.200)	(0.334)	(0.239)	(0.301)
Δpjt	0.435[a]	0.653[a]	0.536[a]	0.399[a]	0.420[a]	0.408[a]	0.327[a]	0.393[a]	0.465[a]
	(0.108)	(0.095)	(0.104)	(0.070)	(0.067)	(0.077)	(0.083)	(0.075)	(0.080)
Constant	0.124	0.028	0.010	0.041[b]	-0.093[a]	-0.089[b]	1.879[a]	1.371[a]	1.334[a]
	(0.121)	(0.036)	(0.025)	(0.018)	(0.025)	(0.035)	(0.379)	(0.202)	(0.232)
Observations	119	119	119	119	119	119	107	107	107
R2	0.216	0.340	0.328	0.326	0.355	0.305	0.378	0.449	0.483

Standard errors in parentheses $^c\, p < 0.10$, $^b\, p < 0.05$, $^a\, p < 0.01$

Conclusion

In this paper, we look at the impact of import restrictions, in the form of tariffs and import bans, on the price of consumption goods in Nigeria. We first develop a simple theoretical model in order to understand the mechanisms for the impact of restrictions between states. The model shows that the price hike associated with a ban differs across states; how it varies with distance to the world market depends on whether the ban effectively chokes off trade. The price hike should increase with distance, provided that smuggled goods do reach the distant states. Turning to the data, we provide graphic and econometric evidence showing that import bans increase the price of goods, with an impact generally smaller in states bordering Benin, a country known to serve for the transit of illegal trade. This is particularly true for agricultural commodities. We then analyze the integration of local markets within Nigeria and find that, for most products under ban, a long-term equilibrium relationship is maintained between prices of goods in remote regional markets. Overall, these results are consistent with the hypothesis that products under ban are being smuggled and reach most states within Nigeria and that states bordering Benin have access to banned products at a lower price, so that smuggling reduces the cost of import prohibitions in these states.

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Appendix

Table 11: Import bans and tariffs

Item	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010
Banana	0	0	0	1	1	1	1	1	0	0
	100	100	100	100	20	20	20	20	20	20
Cabin Biscuit	1	1	1	1	1	1	1	1	0	0
	100	100	100	100	20	20	20	20	20	20
Cerelac-400g	0	0	0	0	1	1	1	1	0	0
	60	60	60	60	10	10	10	10	10	10
Chicken-agric(medium)	0	0	0	1	1	1	1	1	1	1
	35	35	35	35	20	20	20	20	20	20
Detergent-Omo	0	0	0	1	1	1	1	1	1	1
	40	40	40	40	20	20	20	20	20	20
EmbroideryLace-10yds	0	0	0	0	0	1	1	1	0	0
	25	25	25	25	20	20	20	20	20	20
Gari-white	1	1	1	1	1	1	1	1	0	0
	45	45	45	45	20	20	20	20	20	20
Gari-yellow	1	1	1	1	1	1	1	1	0	0
	45	45	45	45	20	20	20	20	20	20
Guinea corn	0	0	0	1	1	1	1	1	0	0
	25	25	25	25	5	5	5	5	5	5
Iodine-bottleof15ml	0	0	0	1	1	1	1	1	1	1
	20	20	20	20	20	20	20	20	20	20
Irish Potato	1	1	1	1	1	1	1	1	0	0
	100	100	100	100	20	20	20	35	35	35
Key soap	0	0	0	1	1	1	1	1	1	1
	100	100	100	100	20	20	20	20	20	20
Maize-grain(white)	0	0	0	1	1	1	1	1	0	0
	25	25	25	25	5	5	5	5	5	5
Maize-grain(yellow)	0	0	0	1	1		1	1	0	0
	25	25	25	25	5		5	5	5	5
Men’s Shoe-Bata	0	0	0	1	1	1	1	1	1	1
	35	35	35	35	20	20	20	20	20	20
Millet	0	0	0	1	1	1	1	1	0	0
	100	100	100	100	5	5	5	5	5	5
Milo-450g	0	0	0	1	1	1	1	1	1	1
	35	35	35	35	20	20	20	20	20	20
Multivite	0	0	0	1	1	1	1	1	1	1
	20	20	20	20	20	20	20	20	20	20
Nivaquine	0	0	0	1	1		1	1	1	1
	20	20	20	20	20		20	20	20	20

Olive Oil-Goya(88.7ml)		0	1	1	1	1				
		65	65	65	50	50				
Orange	0	0	0	1	1	1	1	1	0	0
	100	100	100	100	20	20	20	20	20	20
Palm oil	0	0	1	1	1	1	1	1	1	1
	65	65	65	65	50	50	50	35	35	35
Plantain-ripe	0	0	0	1	1		1	1	0	0
	100	100	100	100	20		20	20	20	20
Plastic Bucket	0	0	0	1	1	1	1	1	1	1
	30	30	30	30	20	20	20	20	20	20
QuakerOats-500g	0	0	0	0	1		1	1	0	0
	60	60	60	60	10		10	10	10	10
Semovita-golden penny	0	0	0	0	1	1	1	1	0	0
	60	60	60	60	10	10	10	10	10	10
Singlet-acrylic(men)	0	0	0	1	1	1	1	1	1	1
	50	50	50	50	20	20	20	20	20	20
Toilet paper	0	0	0	1	1	1	1	1	1	1
	35	35	35	35	20	20	20	20	20	20
Toiletsoap-lux	0	0	0	1	1	1	1	1	1	1
	100	100	100	100	20	20	20	20	20	20
Vegetableoil	0	0	1	1	1	1	1	1	1	1
	20	20	20	20	20	20	20	35	35	35
Women'sShoe-Bata	0	0	0	1	1	1	1	1	1	1
	35	35	35	35	20	20	20	20	20	20

For each item, the first line indicates the presence of an import ban, the second line displays the ad-valorem tariff applying to this good.

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