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The Performance of Bulgarian Food Markets During Reform

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Abstract

Food policy often depends on markets and markets depend on institutions. But how good do institutions have to be before reforms can be launched? Relying on well timed surveys of agricultural prices and a joint study by the Government of Bulgaria and the World Bank on agricultural market institutions, this paper presents evidence that performance in food markets improved following significant policy reforms in Bulgaria, although public institutions remained weak. This suggests that even though strong institutions are preferred to weak ones, it can be costly and impractical to delay policy reforms until work on strengthening institutions is finished. Still, measured performance varied by place and by commodity, suggesting that markets developed at different tempos and that the distribution of benefits from improved markets was uneven. This points to the need to address the costs of adjustment as policies change. The paper introduces a new approach to measure market performance based on composite-error techniques.

This paper—a product of the Sustainable Rural and Urban Development Team, Development Research Group—is part of a larger effort in the department to understand how policies affect food markets in developing countries. Policy Research Working Papers are also posted on the Web at http://econ.worldbank.org. The author may be contacted at DLarson@worldbank.org.

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THE PERFORMANCE OF BULGARIAN FOOD MARKETS DURING REFORM

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Keywords: Bulgaria; composite error; food policy; spatial markets.

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THE PERFORMANCE OF FOOD MARKETS IN TRANSITION: THE CASE OF BULGARIA

1 Introduction

Even though governments intervene in food markets in significant ways, most food policies rely substantively on markets. Within a given policy framework, markets are called upon to signal incentives and to price risk. In addition, market outcomes are often used to measure the success of policy. For many countries this current reliance on markets represents a transition from earlier regimes in which the state took primary responsibility for pricing and sometimes for the physical distribution of food. The focus of this paper, Bulgaria's decision to reduce direct interventions in food markets in favor of a market-reliant system, is a case in point and illustrates a prevalent feature of shifting economic policies during the latter half of the twentieth century.

In the transitional economies of Eastern Europe and countries of the former Soviet Union, the shift to market-based food systems was part of a comprehensive restructuring of national and regional economies. The reforms, and especially the reforms to agriculture, were pursued with varying levels of completeness that reflected uneven political consensus and differences in donor leverage. Nevertheless, more than a decade after the reforms were begun, there remains a commonly expressed view that the costs of adjustment were higher than expected and the benefits of reform slow in arriving because the markets upon which the new policies depended were not up to the task.[1] This criticism relates to the ability of governments to implement reforms, but is more fundamental. Specifically, a reasonable argument is made that, since efficient markets rely heavily on effective publicly backed institutions, the potential benefits of market-based reforms cannot be fully realized until market-supporting institutions can be built up. And, while market reforms often create new incentives and new opportunities, institutions related to property rights, contract enforcement and dispute resolution remain largely the province of governments that are subject to other incentives and competing demands. For this reason, some argue that the pace of reform should be slowed to match an adequate build-up of public institutions.[2]

To make this policy practical, some indication of the readiness of markets is needed. [3] Assessing the capacity of public institutions may not directly reveal the ability of markets to perform however since there is good reason to think that the building up of markets is partly an endogenous

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[1] See, for example, Kydd and Dorward (2001) and references therein.

[2] Ashley and Maxwell (2001) argue that institutions need to be in place before market-dependent policies are introduced. Maxwell (2005) includes a related discussion with reference to the rural development strategies of the European Union, the Food and Agriculture Organization of the United Nations, the International Fund for Agricultural Development and the World Bank

[3] The same could be said for the readiness of institutions. Using examples from Bulgaria, Dimitrova (2002) discusses the difficulties of basing policy on judgments about institutions in the context of EU enlargement.

process in which participants find effective ways to make use of existing limited institutions and, importantly, to build substitutes for public institutions based on private associations and private networks. To the degree that this is so, market reforms that rely on imperfect private substitutes for public institutions may nevertheless create expanded opportunities for trade.

Against this background, our paper takes up the question of whether policy reforms that create new incentives and opportunities can give rise to better performing markets, even when public institutions remain weak. The paper relies on three well timed surveys of agricultural prices in Bulgaria. The first survey took place in 1995 just prior to the 1996 economic crises that would eventually result in a new government and new policies. The second took place in 1997 just prior to a series of agricultural policy reforms that led to the full liberalization of agricultural prices in 1998. The third took place in 2001, following a four-year period when agricultural policies were stable and market-based, but when related institutions remained weak. We also introduce a simple way to measure changes in the performance of spatial markets when information on the underlying determinants of transfer costs is unavailable.

The remainder of the paper is organized as follows. The next section describes events and policies related to food market reforms in Bulgaria during the study period. In section 3, a conceptual model is developed to describe markets that range in their level of completeness and efficiency and an applied version of the model is specified. The data is discussed in section 4. Section 5 contains estimation and simulation results and section 6 concludes.

2 Background

Following the fall of the communist government in 1989, Bulgaria began a process of integrating its economy and institutions with the European Union. Eventually in this process, Bulgarian policy would focus on the transformation of institutions and the building-up of administrative capacity, driven largely by the need to take up and implement the European Union's *acquis communautaire* as a condition of EU membership (Preston 1997; Dimitrova 2002). However, for most of the period between 1989 and the start of negotiations for EU membership in 2000, policy reform in Bulgaria, though uneven, centered on freeing up markets and building private ownership rather than developing institutions. This is particularly evident in the related set of policies that influenced food markets during the seven year period covered in this study.[4]

Political uncertainty marked much of the first decade of reform in Bulgaria and this was reflected in uncertain and inconsistent economic policies. Early on, as the restructuring of the

[4] While the broad criteria for Bulgaria's eventual membership in the EU was laid out by the Copenhagen European Council in 1993, Dimitrova notes that the European Union's emphasis on institutions and administrative capacity began around 1997.

economy began and unemployment rose, many Bulgarians retreated into the agricultural sector, where labor productivity fell and on-farm production strategies shifted to emphasize subsistence. In February 1991, a transitional coalition government launched a broad set of policy changes that included agricultural trade and price liberalization with the intention of boosting food supplies. However, the policies were rendered ineffectual by subsequent interventions, beginning a six-year period of irresolute policy-making. During this time, food policies were short-lived and uncertain. For example, a long-standing policy of fixing prices for basic agricultural commodities was lifted in 1991, but a series of alternative controls were soon adopted in order to slow raising food costs. The controls included minimum procurement prices and recommended retail prices. In 1993 ceiling prices were introduced. In 1995, a State Fund for Agriculture was established and charged with stabilizing prices, although the Fund was given insufficient resources to do so.

By 1996, little progress had been made toward key reforms to Bulgaria's agriculture sector and the state remained a dominant player in the domestic grain market. Private markets were hindered by erratic sales and purchases from the State Grain Reserves, a lack of uniform grades and standards, and an ever changing trade regime characterized by ad hoc export bans thrown up in response to changing food prices (Mergos et al., 2001). Government interventions in related land, water, and credit markets hampered progress as well– a problem that was further exacerbated by the concurrent restructuring of Bulgaria's major trading partners in Eastern Europe and the former Soviet Union.[5]

Following a deepening economic crisis, a grain shortage and widespread bank failures in 1996, a new government was elected to office in 1997. In 1998, agricultural price controls and trade restrictions were lifted as part of a comprehensive reform program. The program was backed by a World Bank loan in 1999, and plans were made for a follow-on loan.[6] In preparation, studies were commissioned on key aspects of the Government's program for agriculture, including strategic grain management, rural finance, land markets, forests, and food safety.

Principally, the studies found that reforms begun in 1998, though substantial, were incomplete.[7] The studies also suggested a need to "move beyond standard liberalization and privatization reforms toward measures to build market institutions" (World Bank 2002, p. 3). A second World Bank loan, approved in June 2001, together with on-going funding from the EU, was

[5] See Deininger (2002) on the general problem of land reform in Eastern Europe; Lerman (2001) discusses the broad problems of agricultural reforms among the transitional countries of Eastern Europe and the former Soviet Union. Mishev, Tzoneva and Ivanova (1998), Csaki, Nash, Fock and Kray (2000), and Mergos, Stoforos, Mishev and Ivanova (2001) review agricultural policies and food markets in Bulgaria.
[6] The two lending operations were: Agricultural Structural Adjustment Loan (PO 57925), approved on June 22, 1999 and Agricultural Structural Adjustment Loan II (PO 57926), approved on June 6, 2001 (World Bank, 2002).
[7] Other sectors important to food markets were subject to inconsistent policies and an incomplete reform process. See, for example, Miller and Petranov (2000).

intended to support already-begun policy reforms and to additionally strengthen a set of institutions judged to be weak. More specifically, the loan backed efforts to strengthen laws and property rights related to agricultural land and forest lands situated on private property. The loan also included provisions design to facilitate the privatization of food processing and related state-owned enterprises, and to put in place legal and regulatory frameworks related to: seed, fertilizer and pesticide markets; inventory-based financing (through warehouse receipts); and irrigation systems.

Transport policies

Coincident with agricultural reforms were a series of policy and investment decisions that affected the transport component of food markets in Bulgaria. In this regard, opening of trade and the freeing of domestic markets during the already mentioned 1991 reforms was crucial because it brought about a reorientation of trade flows and, with it, a restructuring of the Bulgarian transportation sector. This occurred primarily through the emergence of small transport business and external competition rather than changes in basic infrastructure or the privatization of public enterprises.[8] Fuel costs, an important component of transport costs, rose during the study period.[9] Moreover, as with food markets, institutional changes relating to transport markets also occurred late in the reform process.

For example, significant restructuring of the railroad did not take place until the passage of the first Railway Law in November 2001.[10] A second law, passed in early 2002, created a regulatory body and opened the market for transport services. At the same time, rail transport became increasing less relevant to the changing Bulgarian economy and goods carried by rail declined by nearly 61% between 1990 and 2000 (World Bank 2006a; World Bank 2006c). Similarly, despite the emerging importance of road transport, little was done to improve Bulgaria's national road system or the agency charged with its management until late in the reform process. Externally funded projects to improve the road system began around 1994 but centered on the 3,300 kilometers of Class I roads that carried international traffic.[11] For the most part, maintenance efforts were under-funded and proved inadequate. As a consequence, by 2005, only one-third of Bulgaria's national roads were in good condition.[12] Efforts to improve the nearly 4,000 kilometers of regional and district (Class II and

[8] The Bulgarian ground transport network included roughly 4,300 kilometers of rail and 37,000 kilometers of roads in 1990 and in 2001 (World Bank 2005; World Bank 2006a).
[9] By 1992, petroleum prices in Bulgaria were linked to international prices via a cost-plus pricing formula (World Bank, 1992). In constant US dollars, international prices for crude oil rose about 8% between 1995 and 1997, and rose by more than 18% between 1997 and 2001 (World Bank, 2006a). Excise taxes on gasoline, initially low, were raised in 1996 and the phase-out of less expensive leaded gas began in 1998 (Malý, Jakubes, Jilková and Snajdrová, 2002).
[10] The law mandated the separation of the Bulgarian State Railway into two separate enterprises; one in charge of infrastructure and the second in charge of freight and passenger services.
[11] Roughly 19,000 of the 37,000 kilometers of Bulgaria's roads are national roads managed by the Road Executive Agency (REA). The remaining roads are managed locally by cities and towns.
[12] Policies to insure targeted funding for road maintenance were slow to emerge, although a significant step was taken with the establishment of a vignette system in 2004.

III) roads that serve small towns and rural areas were largely delayed until 2006 (World Bank 2005; World Bank 2006b.)

Even though the quality of large portions of the physical transportation infrastructure declined during the study period, several important policy changes did occur between 1997 and 2001 that likely facilitated spatial markets generally. These included an acceleration privatization of transport firms and the development of licensing program that addressed uncertainty over safety.[13]

Policy evaluations

Some germane characteristics of the Bulgarian reform process are captured by the measures given in table 1. The first rows of the table report transition indicators constructed by the European Bank for Reconstruction and Development (EBRD, 2007), while the later rows provide some related indicators of economic performance. The time periods reported correspond to the surveys used in our analysis.

The transition indicators are evaluations of milestones of market reform and range from 1 to 4.33. The privatization scores are average based on the pace of the privatization of small and large enterprises. A score of 2 indicates the development of privatization plans and the beginnings of some sales; a score of 3 indicates that 25 percent of large enterprise assets are sold or ready for sale and that a comprehensive sale of small enterprises in underway. The finance indicator is an average of banking, securities and related reforms. A score of 2 indicates a liberalization of interest rates, the start-up of securities exchanges and rudimentary legal reform. A score of 3 would indicate full interest rate liberalization, significant lending to the private sector, progress on bank regulatory oversight, the emergence of non-bank financial services and an active and regulated securities market. The infrastructure measure is based on reforms covering electric power, railways, roads, telecommunications, water and waste water services. A score of 2 indicates weak commercial objectives and political interference in setting tariffs; some decentralization of decision making. Market reforms are based on progress toward the removal of non-tariff trade restrictions, the elimination of price controls outside of housing transport and natural monopolies, with a score of 4.33 consistent with European Union standards. A competition score of 2 indicates the startup of competition policy legislation and some reduction of entry restrictions; some actions are taken to limit the market power of dominant firms.

[13] As elsewhere in the Bulgarian economy, privatization efforts stalled early in the reform process. In 1997, 233 state-owned enterprises provided freight and passenger service. However, by 1999, all but 23 had been privatized and by 2001 private businesses handled 90% of international freight haulage. Moreover, a series of laws beginning in 1997 introduced EU safety standards, along with inspections and licensing. This reduced uncertainty over the safety of Bulgarian transport companies in the domestic market and, importantly, the international market.

The indicators show that market reforms related to international and domestic trade outpaced reforms in other areas that depended on more specific legislation and the building up of institutions. Still, significant progress was made by 2001 in containing inflation, expanding phone service and improving the overall energy efficiency of the economy – steps that are supportive of domestic markets.

Surveys

During this period of ongoing policy changes, three nationally representative surveys were conducted that asked about food prices. The first survey took place in 1995, prior to the banking crises. The survey was repeated in 1997, just prior to the 1998 reforms and again in April and May of 2001, just prior to an extended period of institutional reform. Of chief interest is the period between the second and third round of surveys, since it is during this interlude that food market reforms were in place but market institutions were judged weak. At the same time, the 1997 survey came on the heels of widespread economic disruption. Consequently, differences in market performance between 1997 and 2001 are potentially attributable to the singular effects of establishing economic stability rather than the related effects of an endogenous development of food markets. For this reason, we make use of the 1995 survey round as well. Together, the three representative surveys allow an evaluation of the combined effects of the set of policy reforms that began in early 1998 against the background of the weak institutions that motivated further lending by the World Bank and the European Union in 2001, and against a 1995 pre-crises benchmark of food market performance.[14]

3 The model

Descriptively, the model starts with the notion that, at any point in time, commodities are stored and offered for sale at spatially diverse locations. The locations are connected by information and distribution systems that collectively comprise a market network. Associated with the network are sets of pair-wise transfer costs relating geographically distinct markets. Some components of overall transaction costs relate to distance or other fixed characteristics of the trading pair. Other components relate to specific characteristics of the traded commodity, such as storability or the ease by which quality can be determined. Other aspects, such as changes in policy, relate to time. Consequently, spatial, temporal and commodity characteristics are expected to play a role in determining observed price spreads. The estimation strategy then is to separate out the components of markets that remain fixed and unaffected by policy from the components that policy can change.

[14] More information about the surveys can be found on the internet at: http://www.worldbank.org/LSMS/ . Annex table 1 provides a chronology of major political events in Bulgaria during the study period.

When competition is sufficient and when trade takes place, the difference between two prices for the same good should differ by no more than the full cost of transacting an exchange between the markets. Consequently, any equilibrium outcome among markets will be characterized by the multiple idiosyncratic relationships between trading partners that reflect location specific storage and transaction costs. From the perspective of an agent operating in a specific location, this solution presents itself in the form of a large number of potential trading relationships, only a few of which are actualized.

Although the general equilibrium outcome among markets at a given point in time is described by the pair-wise price conditions, a variety of factors condition the general multi-market optimization problem. Significant policy reform can change the general problem and consequently give rise to different patterns of trade. In the instance of trade reform, this comes about because trade patterns that are optimal may have been previously precluded by policy. Policy is expected to directly and indirectly affect transfer costs as well since some components of transfer costs are public or club goods. For example, institutional changes, whether induced by changing incentives or supported through policy, are expected to lower transaction costs, especially those related to contracts and information, and in the longer term, to improve the environment for cost-reducing private investments. Similarly, complementary public investment in infrastructure can reduce transfer costs as well.

Even so, more efficient trade does not mean that observed price differences between market locations will necessarily decline or that movements in prices will be more closely correlated. It can be the direct goal of market-constraining policies to minimize price differences among regions and price variation can also be an outcome of monopolistic pricing. Additionally, indirect effects from other market reforms -- for example a reducing in fuel subsidies -- can increase transfer costs and lead to wider price margins.

Nevertheless, food policies are most often incentive-based and reforms are undertaken with the expectation that spatial markets will be less isolated and price transmission across markets will improve. With these policy objectives in mind, the applied model below is designed to measure general indications of these types of change.

Applied model

As discussed, for a given set of initial conditions, a set of trade relationships emerge among trading partners. Though the eventual pattern of trade is determined taking into account the entire set of potential trading relationships, the relationship between any two locations can be described by the

spatial arbitrage inequality.[15] When the storage condition is met, this inequality can be expressed in one of three ways for each market dyad. In the first instance, the following arbitrage equality holds:

$$P_2 - P_1 = T_{12}^* > 0 \tag{3.1}$$

where T_{12}^* is the lowest cost of transfer between locations 1 and 2 and where, in general, $T_{12}^* \neq T_{21}^*$. In the second case, the following condition holds:

$$P_2 - P_1 = T_{12}^* + \alpha_{12} = 0 \tag{3.2}$$

where α_{12} is a positive slack variable. In this case, trade does not take place because demand in location 1 can be met in less costly ways. A third possible outcome is written as $P_2 - P_1 = T_{12} > 0$, where $T_{12} > T_{12}^*$, or equivalently as:

$$P_2 - P_1 = T_{12}^* - \delta_{12} = 0 \tag{3.3}$$

where δ_{12} is also a non-zero slack variable representing missed arbitrage opportunities or constraints that prevent the use of least-cost transfer techniques.[16]

From the perspective of a given set of price dyads (P_2, P_i), where $P_2 > P_i$, the potential relationships can be summarized in terms of a single price-distance equation: $0 < m_{2i} \equiv P_2 - P_i = T_{i2}^* + \alpha_{i2} - \delta_{i2}$ where at least one of the slack variables equals zero at a given point in time, so that $\alpha_{i2}\delta_{i2} = 0$. Were transfer costs easily measured, a solution describing the optimal distribution of supplies across markets and associated slack variables could be found using a variety of programming algorithms. Alternatively, when observations on the determinants of T^* are available, it may be possible to estimate an approximating function.

In the empirical section that follows, we consider price distances relationships across several regions and several commodity markets, and this requires a slightly extended notation. For the purposes of developing the applied model, we denote a particular spatial market relationship by the triad (i, j, c), where i represent the location of a potential exporting community, where j represent the location of a potential importing community, and c represents a commodity. We set up the problem from the perspective of an importing agent who chooses among several potential sources of supply for a particular commodity. Defining market clusters $s(j, c)$, we represent the associated price distance between any two locations at a given time t as $m_{sit} = P_{jct} - P_{ict} > 0$, for all i, j where $i \neq j$.

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[15] Most modern models of spatial integration are derived from Samuelson's (1952) formalization of Enke's (1951) descriptive solution and subsequent extensions by Takayama and Judge (1964).

[16] Constraints related to storage markets can also produce this outcome, although we rule out this possibility by using prices that are observed together with available supplies. See Larson (2007) for a unified discussion of storage and spatial arbitrage conditions. Brennan, Williams and Wright (1997) for a numerical example of how storage and production seasonality can affect trading patterns.

It is worth emphasizing that the subscript s denotes combinations of location and commodity characteristics that remain fixed in the analysis.

With this notation in place and the above discussion in mind, the set of equilibrium outcomes can be written as:

$$m_{sit} = T^*_{st} + \lambda_{sit} \qquad 3.4)$$

where T^*_{st} is the lowest transfer costs for importing regions $j = 1,2,...J$ for commodity $c = 1,2,...C$ for a given set of associated markets $i = 1,2,...I$ and where $\lambda_{sit} = \alpha_{sit} - \delta_{sit}$. In general, transaction costs are not symmetric – for example, in general $T^*_{12c} \neq T^*_{21c}$.

In anticipation of future discussions, note that arbitrage forces will work to minimize δ but not α. Consequently, when all spatial arbitrage opportunities are exploited, λ is non-negative. Moreover, note that when markets become more integrated in the sense that arbitrage takes place among a larger number of market pairings, the number of observations for which $\lambda_{sit} = 0$ increases.[17]

Dealing with missing information on transfer costs

In practice, researchers often have observations on prices at different locations and information on a set of policies expected to affect trade and transfer costs. However, the full set of transportation, transaction and transformation costs that comprise transfer costs are difficult to observe directly and the underlying determinants of transaction costs are observed partially. Consequently, a key component of market integration studies is a strategy to deal with missing information on transfer costs or their determinants. In general this is done by describing the variation in spatial price distances through time. In some cases, the measured price difference is conditional on a partial list of observed determinants – for example transport costs. But most commonly the effects of policy are expected to work through unobserved determinants. The purpose of the modeling effort then is to describe the variation in price distances through time to determine whether outcomes are consistent with the expected policy outcomes. A practical consequence is that most empirical studies of spatial markets are based on the temporal price variation centered on a particular price-pair, or a set of price-pairs.[18]

In terms of the general notation adopted above, the focus of many pair-wise studies is the between-time relationship $m_t = T^*_t + \lambda_t$, where the spatial relationship is fixed and λ is approximated

[17] Put somewhat differently, prices are more often transmitted from one market to the next when the number of pairings with zero-valued slack variables increases.
[18] Fackler and Goodwin (2001) provide a good review of the time-series models and results. See also recent studies by Negassa and Myers (2007) and by Balcombe, Bailey and Brooks (2007) and references therein.

using time-series techniques. More recent models focus on the boundary condition. In this approach, two locations $(1,2)$ are chosen for study based on a priori knowledge about existing trade patterns and two boundaries are identified by approximating transfer costs, $T_{12}(x_t)$ and $T_{21}(x_t)$. Observations that fall outside the boundaries are modeled as $m_{12}(t) = \hat{T}_{12}(t) + \lambda_{12}(t)$ and $m_{21}(t) = \hat{T}_{21}(t) + \lambda_{21}(t)$. A distributional assumption is made regarding the idiosyncratic terms -- normally one in which the $\lambda(t)$ are modeled using a half-normal distribution – which provides a way of estimating the probability that prices remain within the boundaries associated with transfer costs or outside them. In some models, an additional structure is imposed on the $\lambda(t)$ to test whether there are tendencies for prices outside of the two thresholds to move back to threshold levels with time.

Pair-wise threshold models are well suited to studying discontinuities along major trading routes and the seasonality of agricultural markets. However, they are less well suited to studying policy shifts that are expected to remove constraints on trade or significantly affect transfer costs. This is because, as discussed, such policies are likely to engender new trade relationships. In such cases, trade between a given pair may be discontinuous because of events in other markets even when transfer costs are unchanged.[19] Said differently, in times of significant policy shifts, changes in the λ may be partly or fully determined by events in other markets. In this case, the location of the relevant boundary points can shift as can any systemic auto-regressive patterns related to λ.

In this paper, we employ a modified threshold strategy that considers relative price differences among clusters of markets. The approach provides a way to jointly estimate the location of the boundary points and implied transfer costs in multiple markets and provides a measure of how market performance has changed over time. The logic behind the strategy is illustrated in figure 1, which represents the relationship between the equilibrium price in an importing center j and prices in potential exporting centers $i = 1,2,3$. In the first case, the price arbitrage condition holds exactly, and the price distance $p_{jc} - p_{1c}$ exactly equals the associated transfer cost $T_{s(j,c)}$. In the second market, p_{2c} also lies on the transfer cost frontier, but is measured with error e_{2jc}. The third observation lies beyond the frontier by a distance v_{3jc}, which is made up of the slack variable λ_{3jc} and the measurement error e_{3jc}. The estimation problem then is finding a way to separate an estimate of the transfer costs T_s from the idiosyncratic slack variables and measurement errors using the limited information contained in the price data alone.

[19] Samuelson (1952) gives a numerical example.

Statistical specification

With this in mind, first note that when transfer costs are approximated, $T^* = \hat{T}(x) - e$, where e is an estimation error. In this case the estimated boundary condition becomes $m - \hat{T} = v \equiv \lambda + e$. Since neither λ nor e are observed, the distance from the boundary condition can be thought of as a comprised error containing the non-negative idiosyncratic term λ and a random error term e. As it turns out, the components of the comprised error can be separately estimated, based on an assumption about the underlying distributions. In our applied model, we continue to maintain the usual boundary-model assumption that λ can be approximated using the half-normal distribution. This, in addition with the assumption that approximations of the threshold point contain an error e that is distributed normally, results in composite normal-half-normal specification for v.

The amount of information contained in price data alone is limited and it is useful to describe the sources of variation in our panel of markets more precisely. Following Mundlak (1978), let W and B be symmetric and idempotent projection matrixes that generate residuals. For an given vector m of order ST: $Wm = (m_{st} - m_{..})$, $B(s)m = (m_{s.} - m_{..})$, $B(t)m = (m_{.t} - m_{..})$, and $W(st)m = (m_{st} - m_{s.} - m_{.t} + m_{..})$, where the sub-scripted dot represents an average over the relevant index. When the panel is balanced, several orthogonal relationships hold, including:

$$Wm = B(s)m + B(t)m + W(st)m \qquad 3.5)$$

The relationship is axiomatic and exhausts the variation in m. However, the relationships can also be exploited to say something about the distribution of unobserved components, since variation in price distances m are decomposed into average time-invariant effects across markets, $B(s)m$, that are separated from the systemic and idiosyncratic components that change with time, $B(t)m$ and $W(st)m$ respectfully. For the purposes of evaluating policy, the above decomposition distinguishes between differences that have to do with fixed characteristics of the market pairs and those that potentially vary with policy and time.

When all markets are observed for all periods and the panel is balanced, the idiosyncratic within-in terms can be calculated directly. Otherwise, the relationship can be approximated from a regression of m on time and market effects, where the within-market-time effects are recovered as residuals.[20]

$$m_{sit} = \alpha_s + \alpha_t + w_{sit} \qquad 3.6)$$

where α are interpreted as an estimate of the systemic change in transaction costs over markets and over time. The estimates are unbiased as long as the normal OLS assumptions hold. In the context of

[20] In the balanced-panel case, the OLS residuals are identical to the calculated within-market-time effects.

our model, this is most likely when the model is applied to large liquid markets, where the arbitrage condition is expected to hold as an equality.[21] In the general case, when the arbitrage condition sometimes holds as an inequality, it is possible to find better estimates, conditional on an additional assumption about the distribution of the idiosyncratic slack variables.

With this in mind, we estimate a generalized version of 3.6:

$$m_{sit} = \tau_s + \tau_t + v_{sit} \qquad 3.7)$$

where $v \equiv \lambda + e$, and where $e \sim N(0, \sigma_e^2)$ and $\lambda \sim N^+(0, \sigma_\lambda^2)$. Specified in this way, the model can be estimated using stochastic frontier estimation techniques.[22] Note that 3.6 is a special case of 3.7 that holds when $v = e$. Non-parametric programming techniques can also be used to estimate 3.7, although Coelli et al. (2005) discuss reasons why the approach used in this paper is preferred when the $v(t)$ are stochastic.

Caveats

Before proceeding to a description of the data, it is useful to introduce two potential problems that are addressed in the empirical section. The first has to do with the prevalence of opportunities for spatial arbitrage. The modeling approach relies on the assumption that the idiosyncratic slack variables are largely positive in order to separate the idiosyncratic asymmetric λ from the random e. Unexploited arbitrage opportunities result in negative slack variables and if they are frequent in the data, the distinction between the two idiosyncratic terms in the composite error blurs. From the perspective of the estimated model, the issue can be viewed as part of the more general problem of specifying an appropriate composite error. The general problem is much discussed in the stochastic frontier literature and related statistical tests have been devised to resolve the issue empirically. We return to this topic when presenting the empirical results.

The second issue concerns unobserved markets. As discussed, when not all price relationships are observed, there is potential for misidentifying the price relationships that define the trade boundaries. This is an ever-present problem in spatial pricing studies. Making use of clusters of spatial prices arguably makes more extensive use of information about relative pricing. Even so, this does not preclude the possibility that key market relationships are missing. At root is whether or not the data generated by the underlying surveys is sufficiently detailed to support the analysis. This concern cannot be completely resolved; however, as a precaution, we employ bootstrapping methods

[21] See Mundlak and Larson (1992) for a related example.

[22] For early discussions of stochastic frontier models see early Aigner, Lovell and Schmidt (1977), Meeusen and van den Broeck (1977), and by Battese and Cora (1977). Good overviews are given in Kumbhakar and Lovell (2003) and Coelli, Rao, O'Donnell, and Battese (2005).

when generating standard errors for key parameters, which allows us to check whether the results derived from the available data are sensitive to the random exclusion of markets within our sample.

4 Data

As mentioned, prices used in this study were taken from the 1995, 1997 and 2001 Bulgaria Integrated Household Surveys. The 1995 survey included approximately 2,500 households. The second round, in 1997, was conducted using the 1995 sample and approximately 2,000 households were interviewed for both surveys. In 2001, another survey took place that included an identical section on food prices. Because of widespread internal migration, the sample used for the 1997 survey was no longer considered representative. The National Statistical Institute drew a new sample of 2,500 households from the pre-census listing of the 2001 Population Census. Prices on 34 food items, observed in each survey, are used in the analysis.[23] Consequently, the sample constitutes a panel of communities and regions, rather than a panel of households.

Information on all feasible trading partners is included for all spatial markets (i, j, c). In particular, for each of Bulgaria's nine administrative regions, representative prices are derived by finding the modal price for each commodity among the surveyed households of each region. For each commodity, representative prices are then matched to create region-to-region price pairs. Because the prices relate to observed purchases, we assume that associated inventory levels are not so low as to prohibit trade. Under the assumption that transaction costs are non-negative, trade is considered feasible when price distances, measured in logs, are positive. The nine regions yield 72 location dyads for each year and 27 commodities. Not all commodities potentially trade between all dyads. Consequently, the process results in 1,988 observations overall that include 654 observations for 1995, 705 observations for 1997 and 629 observations for 2001.

Table 2 reports average price distances by region and by year. Very generally, the average calculated price distances fall from 1995 to 2001, although not in uniform fashion.[24]

5 Empirical results

As discussed in section 3, the variation in the price distances can be decomposed into between and within components using standard regression or analysis of variation techniques, under the assumption that the within-market-time effects are normally distributed. The composition of the sums-of-squares is reported in table 3. Taken together, the between effects account for more than

[23] The commodities are: beans, beer, bread, bulgar, cabbage, carrots, chicken, cucumbers, fresh milk, ground meat, honey, lamb, lard, lentils, margarine, oranges, pasta, pears, pork, rice, sausages, soft drinks, sweet peas, tomatoes, veal and beef, vegetable oil, and white cheese.

[24] A similar pattern can be observed for price distances based on commodity averages. See table S1 in the supplemental annex.

two-thirds of the overall variation in the sample; however, most of this is from the variation in the between-market component, $B(s)m$, which reflects the variation in fixed location and commodity characteristics. The policies of interest are general in that they apply to all markets in the study, so our primary interest lies with how price differences change with time. The table suggests the change in average values which is captured by the between-time effects $B(t)m$ is small relative to other sources of variation, including the within-time-market residual component, which contains information on how the slack variables vary with time.

Key estimation and test results for the model are reported in table 4.[25] Two of the tests have to do with the chosen specification. The first is a test of the composite-error assumption against the alternative that the errors are symmetrically distributed. The test is due to Schmidt and Lin (1984) and is especially useful since it can serve as the basis of a test of the composite-error assumption that is separate from additional assumptions about the specific form of the composed error. Coelli (1995) proposed a more convenient form of the test, used here, based on the observation that the residuals from an OLS regression will be skewed if the underlying model has a composite-error structure. The test score, reported in the table, takes the expected sign and is significant, lending support assumption that the residuals are a mix of random errors and non-negative slack variables. In the context of the conceptual model, the results are consistent with a distribution of market outcomes containing a mix of the conditions 3.1 and 3.2; the results are inconsistent with a set of markets dominated by missed arbitrage opportunities as described by equation 3.3 or by a set of fully integrated markets as described by 3.1 alone.

The composite composition can be checked in the context of the estimated parameters as well. Note that if the sample contains both positive and negative slack variables in equal share, then the distribution of λ_{st} would appear symmetric and it would be difficult to distinguish the variance of the two components of the composite error. In the context of the estimated model this does not appear to be the case. The estimated natural logs of the variances of λ and e can be recovered directly from the estimated likelihood function and are reported in table 3. The estimated variance of the non-symmetric non-negative portion of the composite error, σ_λ^2, is large relative to the symmetric component σ_e^2. A related test variance-composition test based on the ratio $\sigma_\lambda^2/(\sigma_e^2+\sigma_\lambda^2)$, which goes to zero as the share of total variance attributable to λ declines, is reported in the last line of table 4. For the problem at hand, the statistic, which has a $\chi^2(1)$ distribution, is significant. This is consistent with results from Coelli's test reported above and the

[25] To save space, the large number of fixed-effect estimates associated with the market-commodity dyad, *s(j,c)* are not reported, but are available from the authors.

assumption that the residuals contain a mix of non-negative and normally distributed stochastic terms.

Estimated time-effects associated with the second (1997) and third (2001) surveys are also reported in the table. Because the first-period dummy is suppressed, the time effects can be interpreted as deviations from the 1995 round of surveys. The results suggest that average transfer costs fell between 1995 and 2001. The estimate associated with 1997 is positive, indicating that average transfer costs rose as the economy destabilized; however the associated z-score lies just outside a 10 percent level of significance. In contrast, a test based on the difference between the 1997 and 2002 effects is statistically significant and indicates that average transfer costs fell from 1997 to 2002.[26]

As discussed, a concern in any type of test of the efficiency of spatial markets based on price pairs is that third markets influencing the price outcomes have been omitted from the analysis and that this omission potentially leads to inferential errors. Although we cannot compensate for information about markets that are missing, we test the sensitivity of the results to excluded markets in the context of the sample by randomly excluding some markets and repeating the analysis in a bootstrapping exercise. Results based on 200 replications are reported in table 3. The procedure inflates the standard errors of the respective time-effects, but conclusions based on the first set of standard-errors are not reversed.

The model also provides estimates of the idiosyncratic slack variables, which together with information about relative price changes, can be used provide some insight about differences in individual market outcomes among the three periods. Recall from the earlier discussion that food market reforms of the type undertaken by Bulgaria are expected to lead potentially to different patterns of internal trade and to an improved integration of markets. For any given market pair and commodity, this might mean a reversal of trade flows or possibly a cessation of trade. In terms of the model, this might mean that the inequality $p(1,c) > p(2,c)$ reverses or the value of the associated slack variable, $\lambda(1,2,c)$, might switch from zero to a positive value or both changes might occur together. This makes the equilibrium outcome for any particular market pairing unpredictable. However, in general, market reforms are expected to reduce the disparity in spatial market outcomes, which, in turn, implies that the distribution of the idiosyncratic slack variables should shift toward zero.

[26] Similar results hold for a fixed-effects OLS model, despite strong evidence against in favor of the composite error specification. See supplemental table S2.

As discussed, the sample is unbalanced in that the same triads of commodity and location pairings do not appear in each survey round. This is potentially due to location-specific stock-outs, but is also reasonably attributed to sampling outcomes. The 1,988 observations are based on 939 unique triads. Of these, 299 are observed in all survey rounds.[27] Implied trade reversals are common. Among the 299 repeating triads, the location at which a higher price was observed switch in 124 markets between 1995 and 1997 and in 115 markets between 1997 and 2001, with some markets switching between both periods. Implied unidirectional trade, as indicated by a constant inequality relationship, was a feature of only 124 triads. All of this suggests the potential for shifting trade patterns between the survey periods.

Predicted values for the triad-specific slack variables were recovered from the estimated model. They are calculated as the expected value of the asymmetric component of the composite error, conditional on the jointly estimated symmetric error term, that is $\hat{\lambda}_{st} = E[\lambda_{dt} \mid \hat{e}_{st}]$. Two sample statistics from these predictions are reported in table 5 by survey year for the full sample and also for a sub-sample of the 299 repeating triads. The first is the average value of the slack variable. The second is the average squared value of the slack variable, which corresponds to the average squared distance from the trading frontier.

The predicted slack variables indicate that, once the fixed market effects and the time-difference in transaction costs are taken into account, differences among the survey years in the average value of the slack variables are small. This is true of the overall sample and also of the sub-sample of repeatedly observed markets. The distribution of the slack variables does appear to be more concentrated with time however, as the average squared distance measure falls. This indicates that with time, the distribution of slack variables became more concentrated and more closely bunched near zero. This can also be seen in figure 1, where the density distributions of the estimated slack variables of the repeating triads are graphed by survey year. The results are notionally consistent with improved market integration, as the tail of the distribution attenuates with time.

Taken together, the estimation results indicate that food market performance improved on average as policies implemented in 1998 as average transfer costs fell and as a larger number of markets appeared to integrate.[28] Even so, the results also indicate shifting trade patterns and a significant remaining heterogeneity in outcomes.

[27] Of the remaining unique triads in the sample, 189 are observed once and 451 are observed twice.
[28] The finding complements results from a study by Feinberg and Meurs (2005). Using industry-level data, the authors provide evidence that market reforms in Bulgaria were associated with greater integration between domestic and international markets.

Regional and commodity outcomes

A key attribute of the model specification is that it provides a parametric test of how, in general, markets change over time. This is useful, since the tests provide a way to broadly evaluate the effects of policy. At the same time, the statistical test necessarily compresses information about heterogeneous local outcomes in order to judge a national trend. However, behind general trends heterogeneous local outcomes are expected since the determinants of transfer costs include factors that are specific to commodities and locations.

As discussed, the applied model provides specific predictions for each commodity and market pair at each point in time. Because the market-commodity effects, s, are fixed and the time effects are common to all markets, predicted values are spatially heterogeneous, but rise and fall systematically with time due to the time-effects. However, associated with each observation are idiosyncratic predicted slack variables and these provide an additional component of spatial-temporal heterogeneity.

The degree to which these sources of heterogeneity lead to different outcomes is illustrated in the regional and commodity averages of the predicted price distances given in tables 4 and 5. The predicted transfer costs suggest that the terms of trade between regions and the relative price of traded commodities were affected unevenly by both the disruptions of the late 1990s and the subsequent market reforms. For example, the model indicates that transfer costs fell from period to period in Sofia, albeit at differing rates. In contrast, transfer costs appeared to rise sharply and then decline sharply in Lovetch. The differences may be partly tied to the bundle of underlying commodities observed, since there is large variation in the average predicted transfer costs among commodities. For example, the predicted transfer costs associated with bread remained fairly constant, while many of the predicted transfer costs associated with fresh vegetables changed more dramatically. Collectively, these results illustrate that outcomes from the changes in policy regimes were disparate, even though market performance improved on average in measurable ways.

6 Conclusions

Following widespread economic disorder, the Government of Bulgaria began in 1998 a series of market reforms that included a retreat from previous interventions in domestic food markets. Though the new policies relied heavily on markets, few direct steps were taken to build up the institutions and practices upon which markets depend. Nevertheless, results from the previous section provide evidence that spatial food market performance improved in Bulgaria between 1997 and 2001 following a period of consistent and market-reliant food and transport policies and that food markets functioned better in 2001 than in 1995. Moreover, while the results indicate the

importance of macroeconomic stability for food market performance, there is no evidence to suggest that the observed changes in market margins simply marked a return to earlier performance levels following the economic disruptions of 1996.

Because the improvement in market performance coincided with the introduction of market-reliant reforms and consistent food policy, but took place during a period when related institutions were judged to be weak, the finding is consistent with the notion that some market institutions are endogenous and that, in some cases, there are measurable benefits to consistent policies alone. This speaks against delaying reforms until market-backing public institutions are judged to be strong.

Even so, the estimated changes in market performance are spatially uneven and it may well be the case that this measured heterogeneity is partly related to an uneven evolution of market institutions and related public investments. To the extent this is so, the results suggest that the weak performance of institutions can contribute to an unintended consequences in the distribution of the benefits and costs of reform. Moreover, when privately provided institutions are inferior to public institutions, any potential gains from good policy will be constrained. Consequently, policy makers will need to strike a difficult balance of pacing reforms, building up supporting market institutions and putting in place safety nets where markets are weak.

From a methodological perspective, the results suggest heterogeneous rates of market development, as expressed in the idiosyncratic components of the model. Consequently, methods that impose a fixed dynamic on how prices adjust to policy may produce misleading results. Moreover, the observed pattern of relative price changes and the predicted differences in relative transaction costs are consistent with shifting patterns of internal trade. This is especially important since many studies of spatial markets rely on observations over time of a limited number of markets and on time series approaches where fixed rates of adjustment are assumed.

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Tables

Table 1: Transition and performance indicators.

Transition indicators	**1995**	**1997**	**2001**
Privatization	2.33	2.78	3.22
Finance	2.00	2.34	2.50
Infrastructure	1.33	2.00	2.67
Market reforms	3.34	4.00	4.33
Trade	4.00	4.00	4.33
Domestic prices	2.67	4.00	4.33
Competition	2.00	2.33	2.33
Performance indicators			
Phone lines (per thousand)	307.60	330.97	560.94
Inflation rate	62.05	1,058.37	7.36

Source: EBRD (2007); World Bank (2008).

Table 2: Average price distances by region and year

Region	**1995**	**1997**	**2001**
Bourgass	0.289	0.293	0.259
Haskovo	0.221	0.397	0.187
Lovetch	0.215	0.311	0.196
Montana	0.353	0.310	0.166
Plovdiv	0.308	0.307	0.265
Rousse	0.389	0.306	0.207
Sofia city	0.443	0.345	0.294
Sofia region	0.262	0.286	0.213
Varna	0.318	0.337	0.262
Average	0.321	0.323	0.233
Observations	654	705	629

Source: Bulgaria Integrated Household Surveys and authors' calculations.

Table 3: Analysis of variance for price distances

	Partial sum-of-squares	**Share**
Between effects	80.73[a]	0.69
Time, *B(t)*	2.29[a]	0.02
Markets, *B(s)*	77.33[b]	0.67
Within time-markets, *W(st)*	35.43	0.31
Total	116.16	1.00

[a] and [b] denotes significance at the 1 and 5 percent levels respectively.

Table 4: Selected estimation results and test-scores.

Estimated time effects	**Parameter**	**z-score**	
Estimates		Sample	Bootstrap
τ_{1997}	0.016	1.62	0.75
τ_{2001}	-0.085[a]	-8.00	-3.65
Tests			
Time effects are equal			
$\tau_{1997} = \tau_{2001}$	0.101[a]	9.67	4.29
Slack variables are symmetric			
Coelli skewness test	5.30[a]		
Variance composition test, $\chi 2(1)$	1,324[a]		

*denotes significance at the 1 percent level.

Table 5: Predicted slack variables by survey year

	Full sample		**Repeating markets**	
Survey year	Mean	Squared distance	Mean	Squared distance
1995	0.084	0.054	0.111	0.066
1997	0.076	0.033	0.106	0.044
2001	0.077	0.023	0.107	0.031

Figure 1: Model components

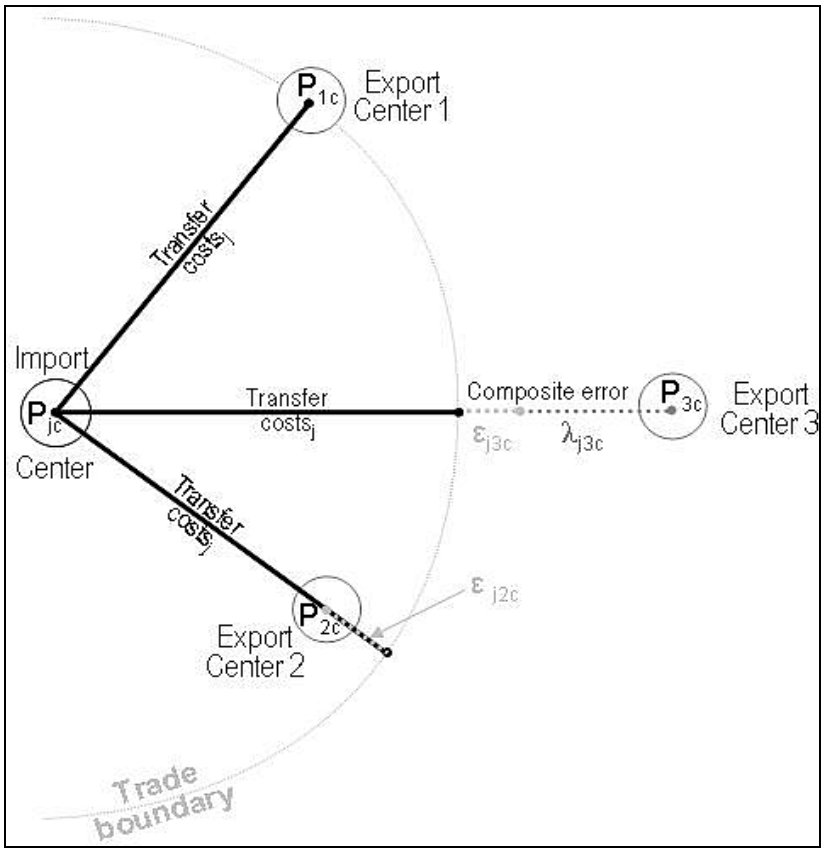


Figure 2: Distribution of estimated slack variables for repeating markets by year of survey

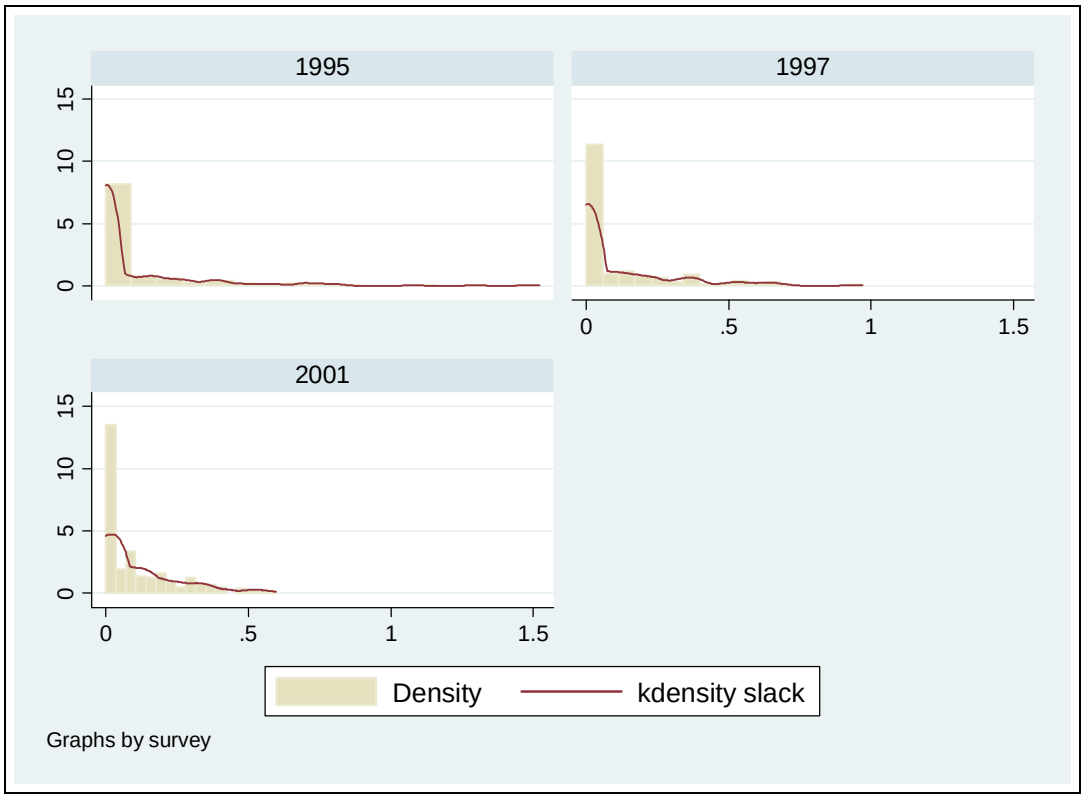


Table 6: Estimated transfer costs by region and survey year.

Region	1995	1997	2001
Bourgass	0.234	0.231	0.140
Haskovo	0.211	0.297	0.141
Lovetch	0.190	0.255	0.162
Montana	0.267	0.274	0.101
Plovdiv	0.234	0.239	0.184
Rousse	0.261	0.232	0.202
Sofia city	0.274	0.252	0.165
Sofia region	0.209	0.216	0.150
Varna	0.227	0.230	0.158
Average	0.236	0.247	0.156

Source: Authors' calculations

Table 7: Estimated transfer costs, by commodity and year

Commodity	1995	1997	2001
beans	0.248	0.204	0.125
beer	0.231	0.293	0.134
bread	0.178	0.169	0.146
bulgar	0.361	0.322	0.189
cabbage	0.309	0.268	0.215
carrots	0.455	0.479	0.213
chicken	0.080	0.171	0.034
cucumbers	0.374	0.408	0.165
fresh milk	0.262	0.282	0.210
ground meat	0.080	0.143	0.005
honey	0.321	0.327	0.253
lamb	0.168	0.226	0.099
lard	0.374	0.427	0.381
lentils	0.261	0.233	0.093
margarine	0.091	0.115	0.021
oranges	0.215	0.283	0.152
pasta	0.199	0.227	0.190
pears	0.128	0.192	0.085
pork	0.082	0.152	0.081
rice	0.117	0.113	0.050
sausages	0.081	0.135	0.032
soft drinks	0.444	0.330	0.482
sweet peas	0.253	0.242	0.190
tomatoes	0.329	0.311	0.095
veal and beef	0.182	0.128	0.126
vegetable oil	0.174	0.183	0.101
white cheese	0.172	0.180	0.107
Average	0.236	0.247	0.156

Note: Authors' calculations

Supplemental tables

Table S1: Average price distances by commodity and year

Commodity	1995	1997	2001
beans	0.333	0.226	0.149
beer	0.261	0.364	0.209
bread	0.218	0.189	0.235
bulgar	0.409	0.364	0.291
cabbage	0.643	0.268	0.229
carrots	0.698	0.792	0.221
chicken	0.087	0.207	0.051
cucumbers	0.480	0.485	0.303
fresh milk	0.262	0.288	0.408
ground meat	0.085	0.204	0.051
honey	0.356	0.355	0.341
lamb	0.193	0.324	0.116
lard	0.445	0.527	0.511
lentils	0.312	0.314	0.105
margarine	0.097	0.124	0.154
oranges	0.219	0.382	0.182
pasta	0.473	0.344	0.235
pears	0.128	0.387	0.105
pork	0.086	0.186	0.180
rice	0.162	0.113	0.113
sausages	0.082	0.215	0.238
soft drinks	0.837	0.408	0.552
sweet peas	0.288	0.302	0.374
tomatoes	0.499	0.428	0.124
veal and beef	0.204	0.148	0.260
vegetable oil	0.182	0.384	0.112
white cheese	0.212	0.182	0.131
Total	0.321	0.323	0.233

Source: Bulgaria Integrated Household Surveys and authors' calculations.

Table S2: Selected OLS estimation results and test-scores.

Estimated time effects	Parameter	Sample t-score	Bootstrap z-score
α_{1997}	0.010	0.56	0.40
α_{2001}	-0.095[a]	-5.36	-4.01
Tests			
Time effects are equal			
$\alpha_{1997} = \alpha_{2001}$	0.105[a]	6.04	5.37

[a] denotes significance at the 1 percent level.

Table S3: Observed markets by year of survey

	Survey year		
Classifications	**1995**	**1997**	**2001**
All markets	654	705	629
Repeating location-commodity combinations	299	299	299
Implied trade direction reversals		124	115
Repeating unidirectional markets	124	124	124