

Holmes, Mark J.; Otero, Jesús; Panagiotidis, Theodore

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On the Dynamics of Gasoline Market Integration in the United States: Evidence from a Pair-wise Approach

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KOÇ UNIVERSITY-TÜSİAD ECONOMIC RESEARCH FORUM
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INTEGRATION IN THE UNITED STATES: EVIDENCE
FROM A PAIR-WISE APPROACH**

Mark J. Holmes
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ON THE DYNAMICS OF GASOLINE MARKET INTEGRATION IN THE UNITED STATES:
EVIDENCE FROM A PAIR-WISE APPROACH ♦

Mark J. Holmes
Department of Economics
Waikato University
New Zealand
holmesmj@waikato.ac.nz

Jesús Otero
Facultad de Economía
Universidad del Rosario
Colombia
jesus.otero@urosario.edu.co

Theodore Panagiotidis
Department of Economics
University of Macedonia
Greece
tpanag@uom.gr

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Abstract

This paper employs a pair-wise approach to examine regional integration in the US gasoline market. Using gasoline price data at the state level over a period of more than two decades, we find strong support for the view that the law of one price holds in regional markets, as more than 80% of bivariate price differentials turn out to be stationary. Furthermore, we uncover evidence that the speed at which prices converge to the long-run equilibrium depends upon the distance between states. Asymmetries are also present in this relationship. Our findings suggest that the more similar are states with respect to taxation, gas stations and refining capacity, the faster is the speed of adjustment towards the long-run equilibrium.

JEL Classification: C33, Q47, R11.

Keywords: Panel data, pair-wise approach, market integration, gasoline, speed of adjustment.

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1. Introduction

Gasoline is a key energy source whose price impact on the economy is well documented; indeed, as indicated by Hamilton (2008), since 1945 nine out of ten of the US recessions were preceded by an increase in oil prices. An examination of regional gasoline price behaviour is important not only from the point of view of understanding the nature and intensity of regional business cycles and consumer prices, but there may also be energy policy implications if markets in different geographic regions have diverged and respond differently to energy price shocks; see Adilov and Samavati (2009). The absence of a long-run relationship between gasoline prices across geographical areas might suggest that a national, or a one size fits all, energy policy for the United States is misguided. If convergence across states is present, then state-level tax policy may be compromised through factors such as cross-border arbitrage. The extent of convergence has further implications for energy policy design. There may be fiscal issues to take into account insofar as relative price dynamics may impact on state budgets, which are in turn influenced by retail gasoline sales. In terms of consumer surplus and welfare, there is an expectation of fair pricing on the part of consumers, where price differentials are wholly attributed to transportation costs, taxation, and other explicit barriers. In addition to this, consumer price convergence can shift an energy tax burden onto the producers, which leads to a competitive disadvantage for firms in high-tax regions (Dreher and Krieger, 2005; Suvankulov et al., 2012).

The retail gasoline sector might at first sight appear to be among one of the most competitive sectors in the US economy, being both deregulated and characterised by disseminated price information in providing a homogenous product where most petrol retailers are independently owned. While such competitive characteristics might be expected to help facilitate closer gasoline price movements across the states, there exist a variety of institutional features that might impede this. Indeed, casual inspection reveals that

the price of gasoline can vary considerably across the US states. The law of one price may not hold across geographically dispersed markets on account of non-zero transportation/transaction constraints on the volume of a good and/or restricted access to the markets (Stigler and Sherwin, 1985; Cuddington and Wang, 2006). Such considerations might inhibit the ability of arbitrage to ensure that prices will be different only to the extent of these costs.

According to Paul et al. (2001), gasoline retailers are not the perfect price takers in a standard economic sense. There is evidence of price discrimination in the retail gasoline market implying some degree of market power. Perfect market integration is also compromised through retail outlets not buying fuel at the same price. Given that fuel is sent to retail outlets by pipeline and trucks from the refineries and marketing terminals, the two main channels for supplying retail outlets are *direct supply* and *jobber distributed*.¹ A factor that is common to both these channels is the presence of vertical integration of gasoline refiners and retailers which can harm competition. In considering price competition, spatial factors have been found to be of importance in pricing decisions by retailers. For example, recent studies such as Iyer and Seetharaman (2008), Lewis (2009), Verlinda (2008) and Bryne (2010) point to the roles played by consumer travel costs, demographics, the degree of local competitive intensity and strategic interaction in influencing the pricing decisions of retail gasoline stations. States vary considerably in terms of the presence and intensity of refining activity. A further institutional consideration concerns local taxation on gasoline which can vary considerably across states. While the average state tax on 1 July 2011 was 22.68 cents per gallon on motor gasoline, state level taxes ranged from 7.5 cents per gallon

¹ As highlighted by Paul et al. (2001), under the direct supply system, retail outlets purchase fuel delivered from the terminals by the refiner. Retail stations supplied under this system may be owned and operated by the refiner or by independent lessee dealers. Alternatively, 'jobber distributed' gasoline is purchased at the terminal by an independent business for the purpose of supplying its own independent stations or for reselling to other dealers. Outlets served in this way may be owned by the refiner, the jobber, or another independent dealer who sells the brand name. In addition, the jobber may also purchase unbranded fuel for distribution to independent stations operating under a non-refiner brand name.

in the case of Georgia up to 37.5 cents in the case of Washington. Such differential tax regimes have the potential to further intensify the heterogeneity in pricing.

Although deregulation of the gasoline market was completed by 1981 (see Paul et al. 2001), there has not been much in the way of a systematic analysis over time of the degree of market integration and price competition across regions of the US gasoline market. Paul et al. (2001) find evidence of a high degree of market integration between prices across five major gasoline markets as evidenced by cointegration tests. However, they conclude that perfect market integration, characterised by a unity slope, is rejected in all but a few cases. In the case of Canadian cities, Suvankulov et al. (2012) find evidence that Canadian retail gasoline markets are well integrated though the share of converging cities reveals a significant decline. Other more recent studies on regional gasoline price behaviour focus on a range of specific issues relevant to regional price variation and adjustment. For example, Lewis (2009) shows that short-lived geographical differences in the severity of wholesale gasoline price spikes are associated with long-lasting geographical differences in retail prices. In a seminal paper, Borestein et al. (1997) demonstrate that gasoline prices rise quickly after an increase in the price of crude oil, but fall slower after a decrease. This asymmetry is questioned later by Bachmeier and Griffin (2003). More recently, Adilov and Samavati (2009) provide results which suggest that gasoline prices could change faster when crude oil prices decrease in some geographic areas. Chouinard and Perloff (2007) note that retail and wholesale gasoline prices vary over time and across geographic locations due to differences in government policies and other factors that affect demand, costs, and market power. Brown et al. (2008) examine wholesale gasoline market integration in the presence of changes in the number of competitors contrasted with geographic market segmentation induced by regulation. Further valuable insights are provided by Deltas (2008), Alm et al. (2009), Kendix and Walls (2010), Marion and Muehlegger (2011) among others.

In this paper we investigate the degree of market integration and test if the law of one

price holds in the long-run for retail gasoline markets after it was completely liberalised in the first half of the 1980s. Cointegration between state gasoline prices is a necessary, but not sufficient, condition for market integration. In order to provide a more direct test of market integration, it is necessary to show that the differences in state prices are stationary or implicitly cointegrated with a unity coefficient. The interaction between non-stationary gasoline price series is analysed drawing on a time-series approach, but in a way that also utilises cross-sectional information. The novelty of our econometric modelling approach is threefold. First, we adopt a pair-wise econometric procedure recently put forward by Pesaran (2007). The idea behind the pair-wise procedure is that for a sample of N states, unit root tests are conducted on all $N(N-1)/2$ price differentials and pairs. The second novelty of our approach is that we employ a range of geographical and economic variables to explain differences in the speed of convergence towards long-run equilibrium across the pair-wise price differentials. In particular, in the event of regional shocks to gasoline prices, we consider the role played by distance between states insofar as the possibility that the speed of adjustment towards long-run equilibrium is fastest between contiguous as opposed to more distant or non-contiguous states. In doing this, we also explore the roles played by the number of state refineries, gasoline stations and tax regimes in influencing the half-life associated with the shock. Third, we consider the role that asymmetries might play in the relationship between speed of adjustment and distance through the application of quantile regression analysis.

Our interest in investigating factors that drive market convergence is in sharp contrast to other studies of the law of one price that focus on different price series and instead consider whether or not regions or countries belong to convergence clusters.² Thus, in relation to the existing literature the innovative contribution of our work is the exploitation of meaningful geographic and economic information to explain differences in

² See, for example, Fischer (2012) and references therein who examine the law of one price in the Euro area.

the degree of convergence towards equilibrium. Clearly, the modelling approach adopted in this paper could be usefully applied to other contexts within and beyond the energy sector as well.³

The paper is organised as follows. Section 2 briefly describes the pair-wise approach for convergence. It is argued that a pair-wise approach towards market integration testing offers important advantages over existing panel data unit root and cointegration methods in terms of addressing the proportion of the sample that is stationary or cointegrated, and the selection of a base or reference state. Section 3 discusses the data and the results of the empirical analysis. Section 4 concludes.

2. Econometric methodology: A brief review

The notion of gasoline price convergence is associated with testing the null hypothesis of a unit root in price differentials. In a sense, support for the alternative hypothesis, that is finding that a gasoline price differential is stationary, is equivalent to saying that the two prices are cointegrated with a known cointegrating vector equal to $[1, -1]'$. Since the unit root tests may include a constant and a deterministic trend term, a rejection of the null implies that regional gasoline prices move together in the long-run but not necessarily such that they are equal.

Our empirical modelling framework adopts the Pesaran (2007) pair-wise testing procedure to analyse probabilistic convergence across a large number of cross section units. In line with Pesaran (2007), we let y_{it} be the observed gasoline price series in state i at time t , where $i = 1, \dots, N$ states and $t = 1, \dots, T$ time observations. The basic idea in Pesaran (2007) is to examine the stationarity properties of all $N(N-1)/2$ possible gasoline market

³ Indeed, some of the examples that can be found in the literature where the pair-wise approach has been applied include the analysis of output convergence (Pesaran, 2007), purchasing power parity (Pesaran et al., 2009), pollutant emissions convergence (Nourry, 2009), and price-level convergence (Yazgan and Yilmazkuday, 2011).

price gaps (or differentials) between states i and j , which we denote as $g_{ijt} = y_{it} - y_{jt}$, where $i = 1, \dots, N-1$ and $j = i+1, \dots, N$. Furthermore, consider the application of the augmented Dickey and Fuller (ADF) (1979) or the Elliott, Rothenberg and Stock (ERS) (1996) unit root tests of order p to the time series $g_{ijt} = y_{it} - y_{jt}$, and let $Z_{ij,T}$ denote an indicator function that is equal to one if the corresponding unit-root test statistic is rejected at significance level α . Thus, for instance, in the case of the ADF unit root test, $Z_{ij,T} = 1$ if $ADF(p) < K_{T,p,\alpha}^{ADF}$, where $ADF(p)$ is the calculated test statistic including p lags of the dependent variable, and $K_{T,p,\alpha}^{ADF}$ is the corresponding critical value for the $ADF(p)$ test of size α , based on T observations. Similarly, when the ERS unit root test is applied, we would have $Z_{ij,T} = 1$ if $ERS(p) < K_{T,p,\alpha}^{ERS}$. Pesaran (2007) studies the fraction of the $N(N-1)/2$ gaps for which the unit-root hypothesis is rejected, and proposes a test statistic given by:

$$\bar{Z}_{NT} = \frac{2}{N(N-1)} \sum_{i=1}^{N-1} \sum_{j=i+1}^N Z_{ij,T}. \quad (1)$$

Under the null hypothesis of non-stationarity, the above statistic has an expected value equal to the nominal size of the underlying unit root test statistic, α . More formally:

$$\lim_{T \rightarrow \infty} E(\bar{Z}_{NT} | H_o) = \alpha. \quad (2)$$

Thus, in the case of a unit-root test such as ADF or ERS, convergence implies that the proportion of rejections is high and should approach 100% as $T \rightarrow \infty$. Analogously, for divergence the proportion of rejections ought to be low and around α . An appealing feature of the pair-wise approach is that it is applicable when the number of cross sections in a panel, N , is large relative to the number of time observations, T .

While pair-wise studies such as Pesaran (2007), Nourry (2009) and Yazgan and Yilmazkuday (2011) focus on computing the fraction of rejections \bar{Z}_{NT} , we progress our

investigation further by calculating the approximated half-life (in months) of a shock for all the stationary price differentials and, more importantly, examining the factor(s) that help explain the speed at which regional prices adjust back towards long-run equilibrium.

An alternative modelling approach based on the estimation of a single vector error correction (VEC) model consisting of the gasoline prices for all the states under consideration would be not feasible, because of the large number of states and lags that would be involved. This approach is employed by Paul et al. (2001) but applied to five Petroleum Administration Defence Districts (PADDs) rather than to state-level data. Of course, there already exist panel unit root tests in the econometrics literature such as Maddala and Wu (1999) and Im et al. (2003), which have been proposed as a means of addressing the low test power associated to univariate methods.⁴ However, Pesaran et al. (2009) observe that the pair-wise methodology offers at least three important advantages over existing panel methods. First, the null hypothesis of these panel unit root tests is that all the series have a unit root, and this joint hypothesis can be rejected even if the proportion of stationary series is small. The pair-wise approach directly addresses the question of what proportion of the gasoline price differentials is stationary. Second, the presence of unobserved common factors complicates the application of panel unit root tests since cross-section dependence leads to size distortions. The so-called second generation panel unit root tests, following the terminology in Breitung and Pesaran (2008), have attempted to allow for possible cross-section dependence through unobserved common factors, but their applications are complicated by the uncertainties surrounding the number of unobserved factors, the nature of the unit root process (whether it is common or individual specific), and the fact that longer data spans are required for taking into account the cross-section dependence. The pair-wise method is robust to cross-sectional dependence. Third, the use of panel unit root tests requires that all series be measured against a common base. In a wider

⁴ See, for example, Fan and Wei (2006), who employ panel unit root test to test the law of one price in China.

sense, this is common practice in studies of regional convergence. However, the outcome of the convergence test can be sensitive to the choice of base region or state.⁵ The pair-wise methodology, by incorporating all the possible bivariate relationships that exist, does not involve what can be a problematic choice of a single reference state in the computation of gasoline price differentials.

3. Data and empirical analysis

We employ monthly price data on regular gasoline sales to end users (measured in dollars per gallon, excluding taxes) for 48 US states. We therefore examine a data set that offers a more comprehensive regional coverage than previous investigations of regional price convergence. The study by Paul et al. (2001) considers all the US states, but for the purpose of analysis these are aggregated into five PADDs each comprising between five and nineteen states. In contrast, our analysis starts off by using state-level data throughout. The price data, which are freely available from the website of the Energy Information Administration (EIA) of the US government at www.eia.gov, cover the period between 1983m1 and 2011m2. However, given that the price data are not available over the period 1987m6 to 1988m12, it was decided to analyse the sample period 1989m1 to 2011m2 which yields a total of 266 time observations for each state. The choice of the sample period is thus dictated by the need to assemble for each state the largest possible uninterrupted price series on regular gasoline sales to end users.⁶ Alaska and Hawaii are excluded from our analysis on the grounds that these states are not geographically contiguous with any other state in the US, so that some of the mechanisms that may underpin long-run constancy of gasoline price

⁵ For example, the gasoline prices of states i and j might be found as non-stationary when measured against a national or base price k , but stationary when measured against one another. This would be the case when there is a highly persistent factor that is common to states i and j , but is not shared by the price k .

⁶ Despite our effort to assemble a balanced panel of data, few sporadic missing values remained: Arkansas (2008m9, 2009m8, 2009m9, 2009m10); Idaho (2010m7); Montana (1996m9, 1999m4, 2001m10, 2002m1, 2002m4); North Dakota (1999m4); New Jersey (2007m1); Nevada (1997m3, 1998m6, 1998m12, 1999m4); Oklahoma (1999m7); Washington (2010m2). These missing values were proxied using spliced values from the local components that would have been used in computing the state measure.

differentials across states within the US may not operate in these cases. The gasoline price series of the 48 states under consideration are plotted in Figure 1, and for the purposes of the empirical analysis are considered in logarithms.

Before implementing the pair-wise approach, it is worth demonstrating the limitations one faces when all series are measured against a common base. Therefore, we apply standard ADF unit-root tests to the relative gasoline price in each state. For this, we (somewhat arbitrarily) select the gasoline prices in four states, namely California (CA), Florida (FL), Illinois (IL) and New York (NY), as alternative base prices with respect to which all other prices will be measured. As can be seen from the results reported in Table 1, although the unit-root hypothesis is rejected most of the time, in some cases the order of integration of the gasoline price depends upon the state with respect it is being measured. For instance, in the cases of the states of Colorado (CO) and South Carolina (SC) we fail to reject the unit-root hypothesis when their prices are measured relative to that in the state of Florida (FL), while for the states of Michigan (MI) and Nevada (NV) failure to reject the null occurs when the prices in these two states are respectively measured against the states of Illinois (IL) and New York (NY).

The next step of our empirical modelling exercise is to calculate the percentage of rejections of the ADF tests based on all 1,128 bivariate gasoline price differentials, see Table 2. Performing the tests on all possible price differentials serves the purpose of avoiding what, in some circumstances, might become a problematic choice of a reference price. The ADF tests are conducted at both the 10% and 5% significance levels, and the order of augmentation of the test regression is determined using the AIC with $p_{\max} = 6$. As can be seen, the percentage of rejections is high, exceeding the size of the unit root test statistics, being equal to 96.01% (92.47%) for the ADF unit root test at the 10% (5%) significance level. Qualitatively similar results are obtained when employing the ERS unit root test.

Clearly, the percentage of rejections is much greater than the corresponding nominal size of the individual unit root tests, so according to our probabilistic definition of convergence we have evidence that is supportive of the law of one price. Our results are consistent with the study by Cuddington and Wang (2006) who employ daily US natural gas spot prices collected at 76 market locations from 1993-1997 and find that 74% of the bilateral price gaps to be stationary. For the stationary natural gas price gaps, the half-lives estimated in this study were in the range of 2 days to 2 weeks.

The strong evidence that the law of one price holds in the long-run suggests that a one size fits all energy policy may not be misguided for the United States. However, our analysis may also have short-run implications for energy policy. Although there is evidence that supports long-run convergence, a varied speed of adjustment towards long-run equilibrium across states might provide some short-run scope for state-level tax policy. It is therefore of interest to consider what factors might drive the speed of adjustment towards long-run equilibrium across states if there is a shock to regional gasoline prices. In order to address this issue, we adopt the methodological framework advocated by Parsley and Wei (1996) in their purchasing power parity study, and use information obtained from the estimated ADF regressions to obtain a measure of the speed of adjustment of the relative price of gasoline between states i and j . This specifically involves employing an approximation of the half-life of a shock to long-run equilibrium based on the estimated autoregressive parameters obtained from the earlier unit root tests. This estimated measure of the half-life of a shock, which we shall refer to as HL_{ij} , is inversely related to the speed of adjustment.

With this consideration in mind, there is an extensive literature that considers the drivers of state-level gasoline prices. For an initial insight into this, we refer to the EIA

analysis of regional petroleum prices⁷, according to which the key considerations are: (i) distance from supply insofar as retail gasoline prices tend to be higher the farther it is sold from the source of supply: ports, refineries, and pipeline and blending terminals; (ii) supply disruptions; (iii) retail competition and operating costs in pump prices; and (iv) environmental programs adding to the cost of production, storage, and distribution.

The earlier introductory discussion highlighted a range of considerations that have been found to impact on the degree of price competitiveness across regional gasoline markets. These include spatial considerations based on consumer travel costs, demographics, the degree of local competitive intensity and strategic interaction, the presence and intensity of refining activity and local taxation. In a related recent investigation, Kendix and Walls (2010) quantify the impact of refinery outages on petroleum product prices and show that refinery outages have a statistically significant positive impact on refined product prices. Marion and Muehlegger (2011) consider several factors that alter the elasticity of supply, including within state heterogeneity in gasoline content requirements, refinery capacity utilization, inventory constraints, and variation in the demand for untaxed uses of diesel. Lewis (2009) studies US gasoline prices following Hurricane Rita and shows that short-lived geographical differences in the severity of wholesale gasoline price spikes are associated with long-lasting geographical differences in retail prices. It is noted that prices may have fallen faster in cities exhibiting retail price cycles. Cycling cities tend to have higher population density and have independent (non-refinery brand) stations that are more highly concentrated into large retail chains. Chouinard and Perloff (2007) note that retail and wholesale gasoline prices vary over time and across geographic locations due to differences in government policies and other factors that affect demand, costs, and market power. Further support is provided by Deltas (2008) who analyses state-level data which suggests that sticky prices and response asymmetries in

⁷ See http://www.eia.gov/energyexplained/index.cfm?page=gasoline_regional.

the gasoline market are, at least partially, a consequence of retail market power, and Alm et al. (2009) who finds that gasoline markets in urban states exhibit full tax shifting onto consumers, but those in rural states (with less competition) demonstrate somewhat less than full shifting.

Following the literature discussed above, one needs to bear in mind the need to obtain consistent data for the key drivers across all 48 states used in our sample. Thus, for each pair-wise HL_{ij} , the state-level drivers that we investigate are as follows. First, we consider a range of cost or supply-side variables that are related to the relative presence of retail outlets, refining activity and taxation across states. These variables include (i) the absolute difference in the logarithm of the number of refineries, $RFRIES_{ij}$; (ii) the absolute difference in the logarithm of crude oil daily processing capacity, $PCPT_{ij}$, where these two variables are included in the analysis on the grounds that they are closely related to an increase in local gasoline supply with less likelihood of incurring fuel transportation costs over long distances; (iii) the absolute difference in the logarithm of the number of gasoline stations, $GASST_{ij}$, since more retail outlets correspond to increased supply and the ability of consumers to shop around; and (iv) the absolute difference in the logarithm of gasoline taxes, $GTAX_{ij}$. While taxes vary across states, the extent of taxes passed onto consumers will depend on how relatively inelastic state demand for gasoline actually is. Second, we consider a range of demand-side variables that includes the absolute difference in the logarithm of the population density, PDS_{ij} , that can be thought of as a driver of the strength in demand. Third, we also include the logarithm of the distance between states, $LDIS_{ij}$. Here, we are particularly interested in examining whether a shorter distance is associated with a faster speed of adjustment back towards long-run equilibrium. Indeed, shorter distances between states may facilitate arbitrage mechanisms that bring fuel prices into line.

Additionally, we create a dummy variable, CB_{ij} , equal to one if two states share a common border, and zero otherwise, and consider its interaction with the supply- and demand-side variables listed above. When referring to these interaction variables, we include the prefix CB ; thus, for instance, $CBGASST_{ij}$ denotes the interaction between CB_{ij} and $GASST_{ij}$, and so on. The appendix provides more detail on the sources and nature of the data.

Using the 1083 pairs for which the unit-root hypothesis is rejected based on the ADF test at the 10% significance level, we initially estimated a regression of HL_{ij} on an intercept, $LDIS_{ij}$, $GASST_{ij}$, PDS_{ij} , $GTAX_{ij}$, $RFRIES_{ij}$ and $PCPT_{ij}$, which also included CB_{ij} along with its interaction with the independent variables already indicated. Perhaps not surprisingly, this initial unrestricted regression was over-parameterised, as several estimated coefficients did not appear either numerically or statistically significant. Thus, after discarding insignificant regressors, we reached the following parsimonious specification:

$$HL_{ij} = -1.067 + 0.536LDIS_{ij} + 1.197CBGTAX_{ij} + 0.574CBGASST_{ij} + 0.018RFRIES_{ij} + \varepsilon_{ij}$$

$$(0.502) \quad (0.075) \quad (0.527) \quad (0.142) \quad (0.007)$$

$$\hat{\sigma}^2 = 1.370$$

where White's heteroskedasticity-consistent standard errors are reported in parentheses (applying the White test for heteroskedasticity of unknown form, including cross product terms, yields an F -statistic equal to 3.129; p -value = 0.000). We first consider the hypothesis that gasoline price relationships between contiguous states might be stronger than between non-contiguous regions. In terms of the pair-wise methodology, statistical evidence of the existence of an positive (inverse) relationship involving distance between any two states and HL_{ij} (the speed of adjustment towards long-run equilibrium) would be consistent with support for this hypothesis. The approximated half-life (in months) and distance (in logs) are plotted in Figure 2. One can observe a clear positive relationship and therefore supportive evidence that is consistent with the hypothesis and spatial effects in

gasoline price convergence. This is supported by a statistically significant coefficient of 0.536 on $LDIS_{ij}$. The estimated coefficients on the other independent variables also turn out to be positive and are statistically significant. This suggests that the more similar are adjacent states with respect to taxation and gas stations, the smaller is the half-life or faster is the speed of adjustment towards long-run equilibrium.⁸ Across all contiguous and non-contiguous states, the more similar is activity with respect to the number of refineries, the faster is the speed of adjustment.

The analysis of the speed of adjustment so far has focused on OLS estimation which provides an average slope coefficient with assumed symmetry in the relationship between distance and speed of adjustment. We further investigate this relationship by considering the possibility that asymmetries may exist. A simple way to explore the latter is by means of quantile regression techniques applied to the above regression; see Koenker and Bassett (1978) and Koenker and Hallock (2001). In comparison to standard linear regression analysis, quantile regression techniques permit us to evaluate the effect of a regressor on the dependent variable not only at the conditional mean of the latter, as in standard linear regression analysis, but also at any particular conditional quantile of its distribution; see e.g. Cameron and Trivedi (2010). This allows for the possibility that the slope coefficient relating the half-life to distance may in fact vary across the quantiles considered. In addition to this, quantile regression techniques offer further attractive features insofar as the resulting estimates are robust in the presence of outlier observations as well as cases where the dependent variable might follow a highly non-normal distribution.

The quantile regression results confirm a positive association on the basis of a median coefficient of 0.540 accompanied by a (Huber Sandwich) standard error of 0.087. Further insight is obtained by examining Figure 3. This reveals that with the exception of the

⁸ Caution must be exercised when interpreting the effect of taxation since taxation levels are evaluated as of 1 July 2011. Ideally one could calculate the average tax over the sample period of interest, so that potential variations in tax levels over a longer time span are accounted for.

0.6 quantile, the slope coefficient relating the half-life to distance is increasing throughout. However, the shaded area indicates that most of the quantiles (from the 0.4 quantile upwards) actually lie within the 95% confidence interval surrounding the OLS coefficient. Overall, these findings confirm a positive relationship between half-life and distance, where the impact from distance has an increasing effect on the half-life starting from the 0.2 quantile. This might be regarded as a threshold effect before which there is no significant slope coefficient. The fact that the estimated coefficient on distance is not statistically different from zero for the 0.1 quantile might be explained by the inclusion of the dummy variable CB_{ij} , that interacts with both $GTAX_{ij}$ and $GASST_{ij}$, which is already picking up the effect of states being adjacent to each other. The slope estimates of the other independent variables included in the model did not appear to vary with the quantile values.

4. Concluding remarks

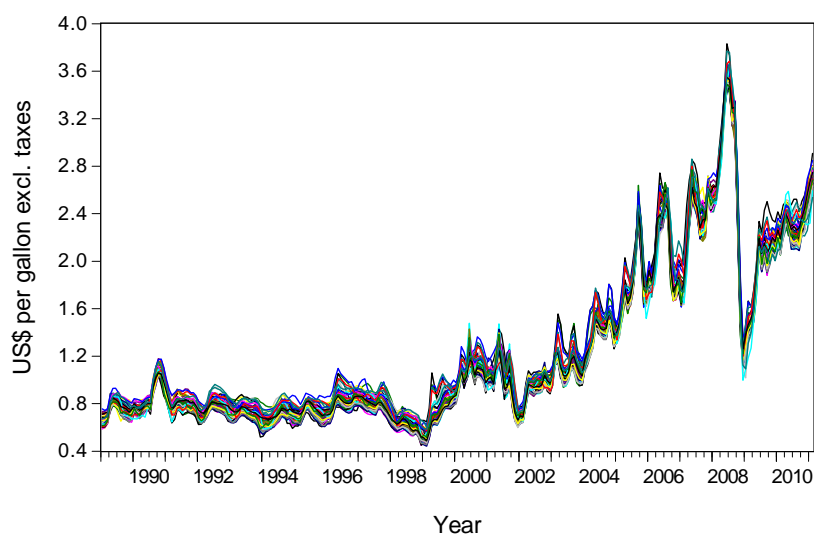
This paper studies integration in the US gasoline market using state-level price data collected over a period of more than two decades. The distinguishing feature of our empirical analysis is that it combines both the time-series with the cross-section dimension. Thus, following Pesaran (2007), we start off by examining the number of gasoline price differentials that can be best characterised as stationary processes. Then, in a subsequent stage of the analysis, we focus on the stationary price differentials and attempt to explain the factors that might help explain the speed at which prices move towards the long-run equilibrium.

We find strong support that the law of one price holds among gasoline markets in the US states, as more than 80% of the gasoline price differentials can be best characterised as stationary processes. Next, focusing on these stationary price differentials, we uncover evidence that the shorter the distance between a pair of states, the less time it takes for the

corresponding prices to adjust. Apart from the geographic separation of markets, our results also reveal that the more similar are states with respect to taxation, gas stations and refining capacity, the faster is the speed of adjustment towards the long-run equilibrium. Our quantile regression analysis suggests that asymmetries may be present insofar as the sensitivity of half-life to distance increases at the lower quantiles.

In terms of energy policy, the support for the law of one price in the long-run suggests that national one size fits all policies are appropriate across the states if one has long-term objectives in mind. However, matters may be different in the short-run. Indeed, our analysis suggests that the speed of adjustment towards long-run equilibrium varies across the state pairs. Furthermore, there are a number of key state-level drivers that influence the half-life of a shock to equilibrium. In this respect, slower speeds of adjustment provide the opportunity for state-level energy policies to have some short-run effect. An interesting policy implication that emerges from our analysis concerns the taxation variable. Indeed, it exemplifies how the speed at which prices adjust to their long-run equilibrium value can be distorted by altering the energy taxation levels of a particular state.

Figure 1. Price on regular gasoline sales to end users in 48 US states



Note: Alaska and Hawaii are excluded from the analysis. The data have been downloaded from the website of the Energy Information Administration of the US government.

Figure 2. Half-life against distance

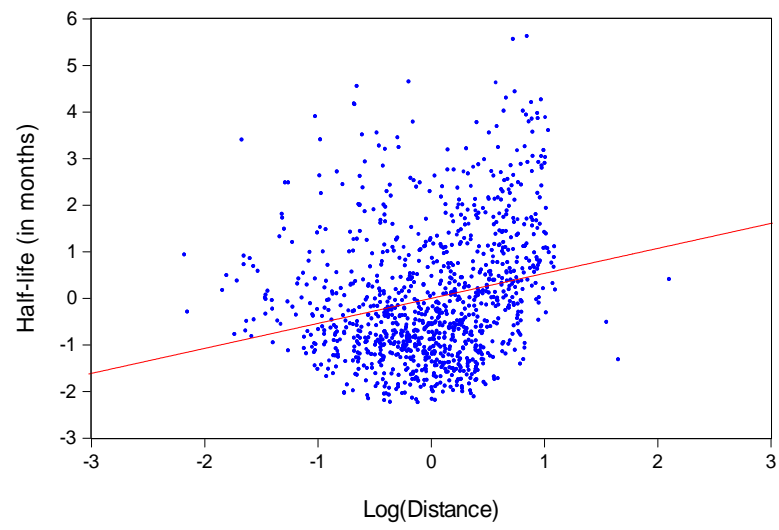
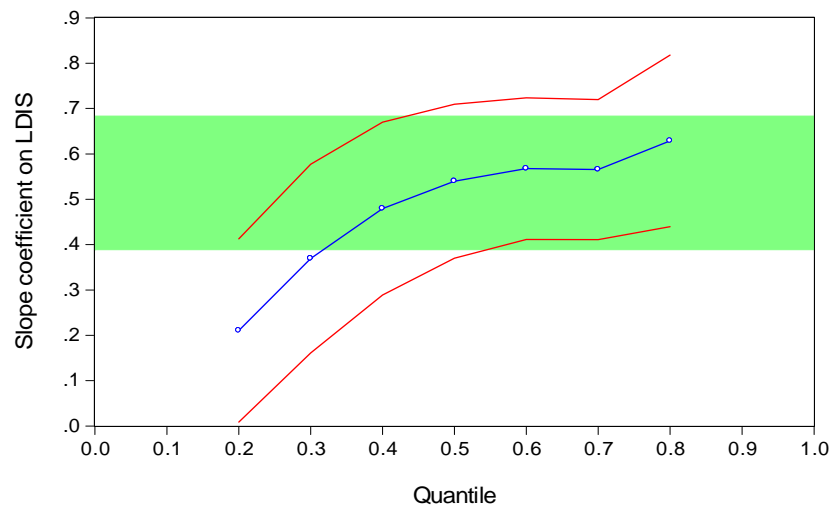


Figure 3. Ordinary least squares and quantile regression estimates



Note: The picture plots the quantile process estimates of the slope coefficient associated to the variable $LDIS_{ij}$ along with the 95% confidence interval. The 95% confidence interval for the OLS slope coefficient (which varies between 0.386 and 0.684) is indicated by the shaded area. Coefficient covariances were calculated using a Huber Sandwich method.

Table 1. ADF unit root t -tests on gasoline relative prices

State	Relative to CA		Relative to FL		Relative to IL		Relative to NY	
AL	-3.554	*	-6.740	*	-3.716	*	-5.286	*
AR	-3.351	*	-5.390	*	-3.185	*	-8.761	*
AZ	-5.054	*	-3.125	*	-4.457	*	-3.692	*
CA	n.a.		-3.875	*	-4.876	*	-4.895	*
CO	-4.175	*	-2.673	*	-4.642	*	-3.693	*
CT	-4.840	*	-7.307	*	-5.565	*	-7.621	*
DE	-4.582	*	-7.071	*	-5.017	*	-7.543	*
FL	-3.875	*	n.a.		-4.194	*	-7.532	*
GA	-4.233	*	-3.569	*	-3.101	*	-5.976	*
IA	-3.634	*	-5.554	*	-2.898	*	-6.951	*
ID	-7.682	*	-3.503	*	-6.873	*	-4.666	*
IL	-4.876	*	-4.194	*	n.a.		-5.796	*
IN	-5.470	*	-5.938	*	-4.002	*	-7.152	*
KS	-3.490	*	-6.965	*	-2.607	*	-7.649	*
KY	-5.135	*	-6.764	*	-5.083	*	-8.225	*
LA	-3.577	*	-4.335	*	-3.733	*	-4.713	*
MA	-4.735	*	-5.338	*	-6.455	*	-7.422	*
MD	-4.285	*	-6.398	*	-4.654	*	-5.832	*
ME	-5.808	*	-6.398	*	-4.277	*	-5.955	*
MI	-4.712	*	-6.212	*	-2.796	*	-7.549	*
MN	-6.526	*	-4.401	*	-4.604	*	-5.288	*
MO	-3.815	*	-6.651	*	-3.787	*	-8.320	*
MS	-4.016	*	-4.502	*	-4.931	*	-8.926	*
MT	-7.172	*	-6.128	*	-4.619	*	-6.635	*
NC	-3.432	*	-3.028	*	-3.719	*	-6.250	*
ND	-4.721	*	-4.799	*	-3.908	*	-4.927	*
NE	-3.662	*	-6.361	*	-2.643	*	-6.618	*
NH	-5.379	*	-5.523	*	-6.559	*	-6.745	*
NJ	-5.984	*	-7.475	*	-6.507	*	-6.959	*
NM	-5.638	*	-5.314	*	-4.559	*	-3.104	*
NV	-5.645	*	-2.897	*	-3.467	*	-2.862	*
NY	-4.895	*	-7.532	*	-5.796	*	n.a.	
OH	-6.136	*	-5.788	*	-7.759	*	-7.235	*
OK	-3.587	*	-7.132	*	-2.435	*	-7.503	*
OR	-5.619	*	-5.521	*	-6.137	*	-4.474	*
PA	-3.499	*	-4.418	*	-3.944	*	-4.920	*
RI	-4.286	*	-6.170	*	-4.915	*	-8.054	*
SC	-3.648	*	-2.807	*	-3.374	*	-5.686	*
SD	-5.179	*	-5.158	*	-3.690	*	-5.514	*
TN	-3.806	*	-6.794	*	-2.693	*	-6.908	*
TX	-2.943	*	-7.050	*	-3.164	*	-7.639	*
UT	-7.514	*	-4.214	*	-6.402	*	-6.578	*
VA	-4.564	*	-7.455	*	-4.890	*	-7.942	*
VT	-4.360	*	-7.637	*	-5.239	*	-6.198	*
WA	-4.245	*	-4.236	*	-6.241	*	-4.672	*
WI	-6.107	*	-5.341	*	-7.497	*	-7.031	*
WV	-3.547	*	-5.943	*	-2.545	*	-5.391	*
WY	-4.696	*	-3.727	*	-6.623	*	-5.393	*

Notes: * denotes significance at the 5% level; n.a. indicates not applicable. See also notes to Table 2.

Table 2. Fraction of rejections when using unit root tests

Significance level	ADF test	ERS test
$\alpha = 0.05$	92.47%	78.81%
$\alpha = 0.10$	96.01%	87.68%

Notes: The ADF unit-root test regressions include a linear trend if it is statistically significant at the 5% level. The number of lags of the dependent variable is determined using the AIC with $p_{\max} = 6$. Critical values for the ADF and ERS tests are based on response surfaces estimated by Cheung and Lai (1995a) and (1995b), respectively.

Data Appendix

In addition to the gasoline price data, which were described in some detail in the main text of the document, we also used data on the following variables:

The distance between states corresponds to the Euclidian distance between the population centres of any two states, based on the geographic coordinates (latitude and longitude) obtained from the Census Bureau for the year 2000, as, used in Garrett, Wagner and Wheelock (2007). The common border dummy variable is based on a zero-one contiguity matrix also used in Garrett, Wagner and Wheelock (2007). We are most grateful to Gary Wagner for kindly providing the distance data as well as the contiguity matrix.

The number of gasoline stations is that on January 31, 2012, as downloaded in the same day from http://www.manta.com/mb_35_B121D7N1_000/filling_stations_gasoline.

State population densities are for the year 2010 as taken from the Current Population Reports of the US Census Bureau.

Gasoline taxes comprise US Federal and State Gasoline Taxes on July 1, 2011 as downloaded on January 31, 2012 from <http://www.commonsensejunction.com/notes/gas-tax-rate.html>.

The number of refineries of each state as well as their crude oil daily processing capacity (in barrels) correspond to data for the year 2010, as taken from the website of the Energy Information Administration of the US government at www.eia.gov.

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