Long-Run Relation among Motor Fuel Use, Vehicle Miles Traveled, Income, and Gas Price for the US

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ABSTRACT

Energy used in transport is a particularly important focus for environment-development studies since such use is increasing in both developed and developing countries and is a carbon-intensive activity everywhere. Gasoline price and per capita motor fuel consumption (and therefore CO_2 emissions) are highly correlated, but it may be too simplistic to assume that higher prices will lead to lower use and thus emissions since there may be a systemic relationship among price, technology, and mobility demand. This paper examines whether a systemic, cointegrated relationship exists among gasoline price, income, and both per capita motor fuel consumption and per capita vehicle-miles traveled, using US yearly data from 1919-2004 and the Johansen cointegration test. The paper's finding of a cointegrating relationship means that gasoline price, technology, and fuel consumption cannot be easily disentangled in the short-run.

JEL classification: Q43

Keywords: Transport energy consumption; Cointegration.

1. Introduction

Energy used in transport is a particularly important focus for environment-development studies since such use is increasing in both developed and developing countries and is (given current technology) a carbon-intensive activity everywhere. (By contrast, for example, electricity generation can be more or less carbon-intensive depending on the energy source used, e.g., coal, natural gas, nuclear, hydro-electric.) Furthermore, understanding the long-run relationship among transport energy use, income, and fuel price in developed countries is important to project with any accuracy global transport fuel use and carbon emissions. Figure 1 shows that vehicle-miles traveled per capita has increased linearly with GDP per capita in the US since 1946. Yet, fuel price may impact vehicle fuel efficiency (miles per gallon), and thus total fuel consumption (more on the relation between fuel price and use bellow).

Figure 1

Economic growth and energy consumption have been highly correlated historically. This correlation coupled with concerns over energy's environmental costs (e.g., carbon emissions) and security issues (e.g., foreign supply dependence and nuclear technology proliferation) has drawn considerable attention to the relationship between energy and development. Some of the literature dealing with this relationship, beginning in the early 1980s, has used statistical techniques from Granger and Sims to reveal the causal direction of the energy-economic growth relationship. More recent advances in the literature have involved improved techniques, like cointegration tests, and updated (and, perhaps, improved) data sets. Taken as a whole, however, the literature on temporal causality between energy consumption and economic growth has offered neither robust results nor convincing rationale.

For example, in the seminal study on the US, Kraft and Kraft (1978) found causality running from GNP to energy consumption for the US over the period 1947-1974. Subsequently, Akarca and Long (1980) shortened the Kraft and Kraft period by two years, while Yu and Hwang (1984) lengthened it by four years, but neither later study detected evidence of causality. A similar history of often contradictory results also emerged for some of the rapidly developing Asian economies. For example, Glasure and Lee (1997) found both cointegration and bi-directional causality for Singapore; however, Masih and Masih (1996), detected neither cointegration nor causality for Singapore. Also, Masih and Masih (1996) found cointegration for both India and Indonesia; however, the direction of causality went from energy to GDP in India, and from GDP to energy in Indonesia. By contrast, Soytas and Sari (2003) detected neither cointegration nor causality for India and Indonesia.

Among the reasons cited for the lack of conclusive or theoretically appealing results are the different data sets, methods for determining lag structure, and statistical techniques (namely, testing for cointegration and whether the analysis considers bi- or multi-variable causality). Still another reason may be the very high level of aggregation of the data analyzed. When considering total energy consumption and total GDP, it is not at all clear what direction causality should be in or how it might evolve temporally. Energy is clearly an input in industrial production; however, in developed countries industry commands a declining share of GDP. Furthermore, a considerable and growing amount of energy consumption in developed countries is for personal transport and use in homes[1]—activities that are “consumptive” in nature, and thus, would be expected to increase with wealth.

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[1] According to Schipper et al. (2001), space heating followed by electrical appliances account for the greatest share of residential end uses in IEA countries.

In an earlier paper, Liddle (2006), I evaluated the link between GDP and energy for a number of OECD countries, using levels of disaggregation of GDP and energy consumption where causation between GDP and energy consumption could be predicted a priori. For example, industry energy consumption—as an input to production—was expected to cause industry GDP, while GDP per capita, or income, was expected to cause energy use in transport and residential buildings—normal consumption goods. The surprising result of the work reported in that paper was that GDP and energy are not strongly linked in most of the countries studied. This paper examines whether a systemic, cointegrated[2] relationship exists among gasoline price, income, and both (i) two definitions of transport demand (per capita motor fuel consumption and per capita vehicle-miles traveled), and (ii) transport technology choice (miles per gallon) over the long-run in the US.

This paper expands on my pervious work in two important ways. First, focusing on transport in the US allows for the consideration of much longer data sets (as long as from 1919 to present) than my, or others', previous work (most analyses are on the order of only 30-40 years). Not having sufficiently long series is a well-known source of spurious results in these types of time series analyses. Second, this paper also considers price and is a multivariate analysis (the previous paper was bivariate, considering only various aggregations of GDP and energy use).

Price is clearly important in determining transport energy use, but a linear causal link between price and such use may not exist. Figure 2 shows a strong relationship between gasoline consumption and gasoline price (an R-squared of 0.48) in IEA countries. However, since these countries are well integrated in the world trade system and gasoline is a commodity, the primary reason gasoline price differs among these countries is taxes. For example, the average pump-

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[2] Two or more nonstationary variables are said to be cointegrated if some linear combination of them is stationary. The finding of cointegration among economic variables is interpreted as proof of a long-run, equilibrium

price of gasoline (in USD/liter) for the largest eight economies in the OECD was 0.93, during March 2003; the standard deviation was 0.32, and the range [0.42, 1.23]. However, excluding taxes the average price, standard deviation, and range were 0.34, 0.04, and [0.29, 0.42], respectively (data from the International Energy Agency). The ability of (democratic) governments to tax gasoline (and thus, personal transport), must to some degree, reflect relative demand for mobility. In other words, since gasoline price does influence technology choice and mobility demand, and, since gasoline price is heavily influenced by government in both market and non-market economies, technology and mobility demand likely affect price too. Of course, given a particular need for mobility, a higher fuel price might encourage a more fuel-efficient vehicle fleet, and thus lower fuel consumption.

Figure 2

The following section of the paper introduces the data and methodology used. Section 3 presents and discusses the results. The final section summarizes the conclusions and addresses some policy implications.

2. Data and Methodology

The data series used in this study are annual data, converted to natural logs. The series are: real per capita GDP, 1919-2004 from Johnston and Williamson (2005); vehicle-miles traveled per capita, 1936-2004, and motor fuel use per capita, 1919-2004, both from the US Department of Transportation, Federal Highway Administration’s Highway Statistics; miles per gallon, 1936-2004 (vehicle-miles traveled divided by motor fuel use); and real retail gasoline price from US Department of Energy, Energy Information Agency. Population data (to convert measures to per capita) is from the US Census Bureau. Table 1 summarizes the data series.

Table 1

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relationship. More on cointegration follows in the Data and Methodology section.

The first step is to test for unit roots in each series since all variables in a cointegration test should be of the same order. It is expected, as others have found, that these series (all of which contain noticeable trends) will be nonstationary in levels, but stationary in first differences. To test for unit roots, I use both the Elliott, Rothenberg, and Stock (1996) Dickey-Fuller test with GLS detrending (DF-GLS), and the Ng-Perron test (Ng and Perron, 2001) also with GLS detrending. These tests are particularly appropriate for highly trending data; furthermore, Maddala and Kim (2000) argue that DF-GLS tests are more powerful than the (often used) augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. Both tests allow for a constant or a constant and a linear time trend in the test regression. For robustness, I report the results from both types of equations and both tests in Appendix Tables A1-a and A1-b. The power of unit root tests is sensitive to the number of lagged terms used. To choose the optimal number of lags, for augmented Dickey-Fuller tests, I employ Hall's (1994) *general to specific rule*, where one starts with a maximum number of lags, tests the significance of coefficient on the last lagged term, and reduces the number of lags iteratively until a significant statistic is encountered. For the Ng-Perron tests, the number of lags is determined by the Schwartz information criterion.

Engle and Granger (1987) pointed out that a linear combination of two or more nonstationary series may be stationary. If such a stationary linear combination exists, the nonstationary time series are said to be cointegrated. The stationary linear combination is called the cointegrating equation and may be interpreted as a long-run equilibrium relationship among the variables. The purpose of the cointegration test is to determine whether a group of non-stationary series are cointegrated or not. To test for cointegration, I use Johansen and Juselius' (1990) test for multivariate cointegration. The Johansen cointegration approach produces two statistics (the trace and maximum eigenvalue statistics), which can conflict—although they do not in the results presented here. To determine the number of cointegrating relations, *r*, one proceeds sequentially from $r = 0$ to $r = k - 1$, where *k* is the number of endogenous variables, until one fails to reject. The two tests of the null hypothesis of *r* cointegrating relations against the alternative of *k* cointegrating relations, for r = 0, 1, …, *k* - 1 are reported in the results tables.

3. Results and Discussion

The results of the unit root tests using the modified Dickey-Fuller test appear in Appendix Table A1-a, and the results of the unit root tests using the Ng-Perron test are shown in Appendix Table A1-b. Different unit root tests sometimes produce inconsistent results. However, since I expected (as others have found) all of these series to be order I (1), i.e., stationary in first differences but not in levels, I interpret the results shown in Tables A1-a & A1-b not to show convincing evidence to reject my prior belief that all of these series are I (1).

Three sets of cointegration tests were run: one involving each definition of mobility demand (motor fuel use and vehicle-miles traveled) and one involving technology choice (miles per gallon). Tables 2 and 3 show that, for the two sets of transport demand variables—GDP per capita, price, and motor fuel use in Table 2, and GDP per capita, price, and vehicle-miles traveled in Table 3—both the trace and max-eigenvalue test statistics indicate one cointegrating equation at the one percent significance level. Table 4 shows that, for technology choice (miles per gallon), price, and GDP per capita, both the trace and max-eigenvalue test statistics indicate one cointegrating equation at the five percent significance level. These findings of cointegration confirm a long-run, systemic relationship among price, income, and transport demand or technology choice in the US.

Tables 2 - 4

4. Summary, Conclusions, and Policy Implications

Transport is a major consumer of energy and an important source of carbon dioxide emissions everywhere. The demand for mobility increases strongly with income. Although gasoline price and gasoline consumption are correlated across countries, it may be too simplistic to assume that higher gasoline taxes in high consuming countries would result in much lower fuel consumption, and thus carbon emissions. The analysis here showed that in the US mobility demand (proxied by either motor fuel use or vehicle-miles traveled) has a long-run systemic relationship with gasoline price and income, as does technology choice (miles per gallon) with gasoline price and income. And thus, these variables cannot be easily disentangled in the short-run. This paper's finding of a cointegrating relationship has two important implications for policy. First, in countries like the US, higher standards for vehicle efficiency are preferable to higher gasoline taxes since (i) the former may lead to a faster change in overall vehicle fleet efficiency; and (ii) the level of taxes necessary for the latter approach to cause a sufficient change in vehicle fleet efficiency via the market would be quite painful. Second, countries like China and India would be wise to develop a system of prices, technology, and mobility options that help them avoid the difficult choice that the US now faces.

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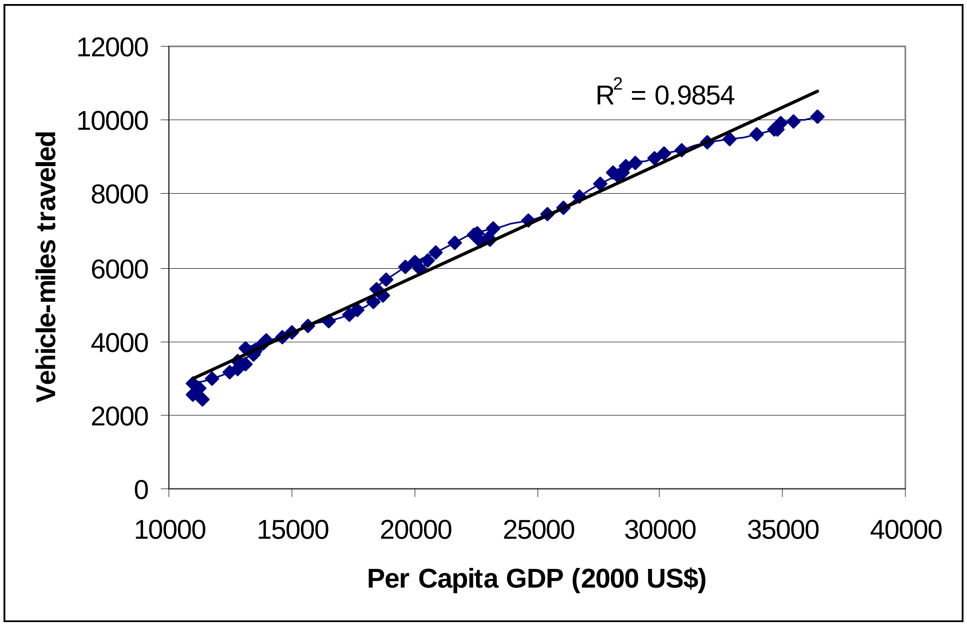

Figure 1. Vehicle-miles traveled per capita and GDP per capita in the US, 1946-2004. GDP data are from Johnston and Williamson (2005). Travel data are from the Federal Highway Administration.

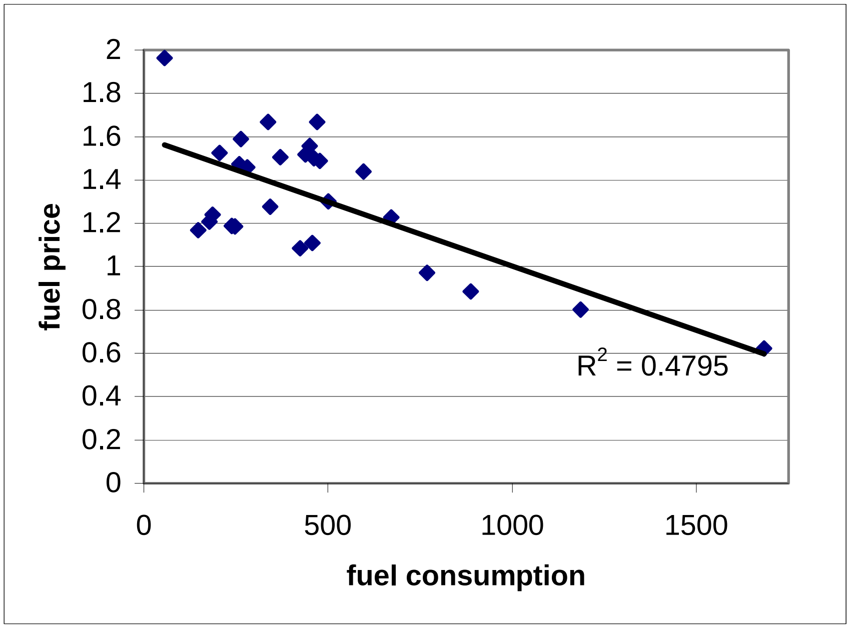

Figure 2. The relationship between gasoline consumption (in liters per capita) and unleaded gasoline price (in US dollars per liter) for IEA countries. Data is from the International Energy Agency 2005 & 2006.

Table 1. Variable definitions and data sources.

Variable	Definition	Observations	Source
LGDP	Natural log real GDP per capita	Annual, 1919-2004	Johnston and Williamson (2005) http://www.eh.net/hmit/gdp/
LPRICE	Natural log real retail gasoline price	Annual, 1919-2004	Energy Information Agency http://www.eia.doe.gov/emeu/steo/pub/fsheets/PetroleumPrices_files/frame.htm
LMFU	Natural log motor fuel use per capita	Annual, 1919-2004	Federal Highway Administration http://www.fhwa.dot.gov/policy/ohpi/hss/hsspubs.htm
LVMT	Natural log vehicle-miles traveled per capita	Annual, 1936-2004	Federal Highway Administration http://www.fhwa.dot.gov/policy/ohpi/hss/hsspubs.htm
LMPG	Natural log of miles per gallon, calculated as vehicle-miles traveled divided by motor fuel use	Annual, 1936-2004	See above

Table 2. Johansen and Juselius cointegration test for GDP per capita, prices, and motor fuel use, 1919-2004

Null	Alternative	Statistic	Critical values	
			5 %	1 %
Variables: LGDP, LPRICE, LMFU				
Trace statistic				
r = 0	r > 0	37.08**	29.68	35.65
r <= 1	r > 1	9.61	15.41	20.04
r <= 2	r = 3	2.13	3.76	6.65
Maximum eigenvalues				
r = 0	r > 0	27.46**	20.97	25.52
r <= 1	r > 1	7.48	14.07	18.63
r <= 2	r = 3	2.13	3.76	6.65

r: Indicates the number of cointegrating relationships. Trace test and maximum eigenvalue statistics are compared with the critical values from Johansen and Juselius (1990). ** Indicates rejection of the null hypothesis of no cointegration at the 1% level.

Table 3. Johansen and Juselius cointegration test for GDP per capita, prices, and vehicle-miles traveled, 1936-2004

Null	Alternative	Statistic	Critical values	
			5 %	1 %
Variables: LGDP, LPRICE, LVMT				
Trace statistic				
r = 0	r > 0	39.00**	29.68	35.65
r <= 1	r > 1	13.23	15.41	20.04
r <= 2	r = 3	4.42	3.76	6.65
Maximum eigenvalues				
r = 0	r > 0	25.77**	20.97	25.52
r <= 1	r > 1	8.81	14.07	18.63
r <= 2	r = 3	4.42	3.76	6.65

r: Indicates the number of cointegrating relationships. Trace test and maximum eigenvalue statistics are compared with the critical values from Johansen and Juselius (1990). ** Indicates rejection of the null hypothesis of no cointegration at the 1% level.

Table 4. Johansen and Juselius cointegration test for GDP per capita, prices, and miles per gallon, 1936-2004

Null	Alternative	Statistic	Critical values	
			5 %	1 %
Variables: LGDP, LPRICE, LMPG				
Trace statistic				
r = 0	r > 0	32.69*	29.68	35.65
r <= 1	r > 1	7.90	15.41	20.04
r <= 2	r = 3	0.05	3.76	6.65
Maximum eigenvalues				
r = 0	r > 0	24.78*	20.97	25.52
r <= 1	r > 1	7.85	14.07	18.63
r <= 2	r = 3	0.05	3.76	6.65

r: Indicates the number of cointegrating relationships. Trace test and maximum eigenvalue statistics are compared with the critical values from Johansen and Juselius (1990). * Indicates rejection of the null hypothesis of no cointegration at the 5% level.

Table A1-a

Results from unit root tests on levels and first differences of GDP per capita, prices, motor fuel use, vehicle-miles traveled, and miles per gallon using Dickey-Fuller with GLS detrending

	Levels		First differences	
	Trend & constant	Constant	Trend & constant	Constant
LGDP	-3.80 [1]***	1.03 [11]	-2.24 [10]	-0.97 [10]
LPRICE	-3.26 [4]*	-0.92 [2]	-7.12 [1]***	-6.99 [1]***
LMFU	-0.91 [5]	0.77 [5]	-1.73 [4]	-0.50 [4]
LVMT	-2.14 [4]	0.84 [5]	-5.10 [0]***	-4.81 [0]***
LMPG	-1.24 [0]	-0.50 [0]	-6.09 [0]***	-2.67 [0]***

The Elliott-Rothenberg-Stock DF-GLS test statistic is shown. The numbers in brackets are the optimal lags determined by Hall’s general-to-specific procedure. Levels of significance are indicated by ***, **, and *, referring to the 1%, 5%, and 10% levels, respectively.

Table A1-b

Results from unit root tests on levels and first differences of GDP per capita, prices, motor fuel use, vehicle-miles traveled, and miles per gallon using Ng and Perron test with GLS detrending

	Levels				First differences			
	Trend & constant		Constant		Trend & constant		Constant	
	Z_a	Z_t	Z_a	Z_t	Z_a	Z_t	Z_a	Z_t
LGDP	-23.37 [1]**	-3.41 [1]**	0.75 [1]	0.54 [1]	-17.87 [0]**	-2.98 [0]**	-10.07 [0]**	-2.20 [0]**
LPRICE	-22.76 [1]**	-3.03 [1]**	-2.83 [0]	-1.11 [0]	-41.31 [0]***	-4.40 [0]***	-40.11 [0]***	-4.40 [0]***
LMFU	0.63 [3]	0.80 [3]	0.62 [3]	0.80 [3]	-22.70 [0]**	-3.34 [0]**	-0.92 [4]	-0.55 [4]
LVMT	0.95 [4]	0.97 [4]	0.95 [4]	0.97 [4]	-27.31 [0]***	-3.69 [0]***	-25.40 [0]***	-3.56 [0]***
LMPG	-2.29 [0]	-1.02 [0]	0.70 [0]	0.53 [0]	-31.34 [0]***	-3.95 [0]***	-12.69 [0]**	-2.52 [0]**

The Ng-Perron test statistics are shown. The numbers in brackets are the optimal lags determined by the Schwartz information criterion. Levels of significance are indicated by ***, **, and *, referring to the 1%, 5%, and 10% levels, respectively.